Growth, Financial Development, Market Liquidity and Risk

A thesis submitted for the degree of Doctor of Philosophy

by

Bin Tan

Economics and Finance, School of Social Science

Brunel University

London, UK

DECEMBER 2010

Abstract

This thesis, firstly, studies the impact of financial liberalization and political instability on economic growth and quantitatively examines the relative importance of the identified underling reasons of Argentine riddle by using an innovative econometric methodology and unique data set: it presents power ARCH estimates for Argentina from 1896 to 2000. The main results show that the long-run effect of financial liberalization on economic growth is positive while the short-run effect is negative, albeit substantially smaller. The political instability effects are substantially larger in the short-run than in the long-run. We also investigate potential mechanisms for the effects of financial liberalization and political instability on economic growth: direct impact or happening through the variation of growth volatility. Our results also suggest that financial development, trade openness and political instability are the main factors to explain the Argentine decline.

Furthermore, real business cycle variability - growth relationship and the link between inflation and its uncertainty are investigated by using monthly data of four Asian countries/regions (Japan, South Korea, Singapore and Taiwan) and parametric power ARCH methodology to proxy uncertainty. We find that more uncertainty about output leads to a higher rate of growth in three of the four countries/regions and the form of the uncertainty matters. Output growth reduces its uncertainty in all countries/regions via inflation uncertainty except Singapore. For all countries/regions, inflation significantly raises inflation uncertainty as predicted by Friedman. On the other hand, increased uncertainty affects inflation positively in Japan and Singapore, which support the Cukierman-Meltzer hypothesis. We find a negative sign for Taiwan which is in accordance with the Holland hypothesis when error term was normally distributed, however, this result is not statistically significant when the student-t distribution is applied. Interestingly, South Korea's data reveals a positive sign initially, however, it turns around when a structural dummy is incorporated. This dramatic outcome in favour of the Holland hypothesis, and chimes in with Dueker and Kim (1999), who claim that the inflation was strictly controlled by the South Korean monetary authority.

In addition, this thesis investigates two-way causal relationships between spread, volatility and volume in the FTSE100 stock index over the period from 1992 to 2004 by using bivariate AR-FI-GARCH model and multiple measurements of risk and spread. The measurements of the spread include relative bid-ask spread, effective bid-ask spread, the inventory cost component of the bid-ask spread and the information cost component of the bid-ask spread. Risk is proxied by two measurements of price volatility: the close-to-close volatility and the range-based volatility. We also take the impact of electronic trading into account. Our results suggest that the spread and volume are positively impacted by volatility simultaneously. In addition, both volatility and volume are negatively affected by the spread. Furthermore, we find that the inventory cost component of the spread has a negative effect on volatility, in contrast, the information component of the spread positively impacts volatility. These results support the argument that speculation generates volatility in the market and higher transaction costs benefit stability of the market.

Acknowledgments

I would like to express my sincere appreciation to my supervisor Professor Menelaos Karanasos for his thoughtful guidance, effortful help and inspiring encouragement. I also want to thank my coauthors Professor Nauro Campos and Professor Andros Gregoriou for their invaluable help and support.

Many thanks are due to PhD students and faculty of Economics and Finance Department of Brunel University. I really enjoyed our seminars, conferences and discussions. It was a great pleasure to study in this friendly collectivity.

Finally, I want to take this opportunity to thank my parents who always give me support and encouragement.

Declaration

Chapters 1 and 2 were written in collaboration with Professor Menelaos Karanasos and Professor Nauro Campos. Chapter 1 has been sent to the Journal of Banking and Finance and we have been invited to revise and resubmit it.

Chapters 3 and 4 are joint works with Professor Menelaos Karanasos.

Chapter 5 was written jointly with Professor Menelaos Karanasos and Professor Andros Gregoriou.

Contents

List of Tables

List of Figures

Introduction

Instability and performance are often inversely related. Financial crises are associated with growth decelerations and contractions, while political protest tends to disrupt productive activities thereby negatively affecting economic growth. Such amplified uncertainties, driven either by economic or political events, have deleterious consequences in terms of economic performance, especially in the short-run. In the long-run, however, financial development and political instability may instead have positive effects on growth. For example, the supply of credit to the private sector and transitions from autocracy to democracy are often considered as key determinants of long-run growth across countries. Seminal papers are those by Kaminsky and Schmukler (2003), Tornell et al. (2004) and Loayza and Rancière (2006) argue that despite the development of the financial system being robustly associated with economic growth, it is also often found to be the main predictor of financial crises. That is, while the long-run effect of finance on growth is positive, in the short-run it is negative. However, cross-country heterogeneity and business cycles synchronization issues may play an undesirably large role in generating this result, in general, and in particular regarding the relative magnitudes of these two effects. For instance, Loayza and Rancière (2006) report that the size of the effects is similar but the negative short-run effect is often larger than the positive long-run effect. To put forward a deeper understanding of these relationships we try to answer the following questions. What is the relation between financial development on the one hand and economic growth and its volatility on the other? Does the sign and intensity of such effects vary over time and do they vary with respect to short- versus long-run considerations? Is there a dynamic asymmetry in the impact of financial development (that is, is it negative in the short- and positive in the long-run)?

Besides the financial development, the fluctuation of economic performance/variability in the business cycle affects economic growth as well. Until the early 1980s, macroeconomic theorists treated the analysis of the real business cycle (RBC) as separate from the study of economic growth. In the 1980s, three important contributions in business cycle theory by Kydland and Prescott (1982), Long and Plosser (1983), and King et al. (1988) integrated the theories of the business cycle and economic growth in their models. However, these models did not consider the possibility that the variability of the business cycle might relate to the rate of economic growth. Similarly, for the most part, developments in growth theory have been made without consideration of the variability in the business cycle. The scene has changed recently at both the theoretical and empirical front. At the theoretical level, Blackburn and Pelloni (2005) and a number of studies summarised by these authors examine how cyclical fluctuations might relate to long-run economic growth. At the empirical level, highlight the importance of the reduction in US GDP growth volatility in the last two decades and its implications for growth theory. The early dichotomy in macroeconomic theory between economic growth and the variability of economic fluctuations should be reconsidered given several theories regarding the relationship between output volatility and growth predict a positive, negative or no association between the two variables. The empirical evidence to date based on cross-section country studies, panel data studies, or time-series analyses of individual countries are also quite mixed. The theoretical and empirical ambiguity surrounding the RBC variability–economic growth relationship provides us with the motivation to expand on the empirical aspects of this issue.

In order to promote economic growth, many countries (especially Asian countries such as South Korea) adopt export oriented growth strategy which is normally associated with high inflation. The issue of the welfare costs of inflation has been one of the most researched topics in macroeconomics both on the theoretical and empirical fronts. Friedman (1977) argues that a rise in inflation leads to more nominal uncertainty. The opposite direction of causation has also been analyzed in the theoretical literature. Cukierman and Meltzer (1986) argue that central banks tend to create inflation surprises in the presence of more nominal uncertainty. Clarida et al. (1999) emphasize the fact that since the late 1980s a stream of empirical work has presented evidence that monetary policy may have important effects on real activity. Consequently, there has been a great resurgence of interest in the issue of how to conduct monetary policy. If an increase in the rate of inflation causes an increase in its uncertainty, one can conclude that greater uncertainty which many have found to be negatively correlated to economic activity is part of the costs of inflation. Thus, if we attempt to provide a satisfactory answer to the questions "What actions should central bankers take?" and "What is the optimal strategy for monetary authorities to follow?" we must first develop a clear view about the temporal ordering of inflation and nominal uncertainty.

Steady economic growth needs support of a healthy financial market which has high liquidity and low fluctuation. One of the most critical factors that investors look for in any financial market is liquidity. Liquidity is defined as the ability to trade stock rapidly with little price impact. To maintain liquidity, exchanges use market makers, who are individuals willing to provide a financial market whenever the investors wish to trade. In return for providing the financial market, market makers are granted monopoly rights by the exchange to post different prices for stock purchases and sales. As a result, market makers buy stock at the bid price and sell the stock at the higher ask price. This ability to buy the stock low and sell high is the market makers' compensation for providing liquidity in the financial market. Their compensation is defined as the ask price minus the bid price, which in turn is defined as the bid-ask spread. Studies such as Atkins and Dyl (1997), Constantinides (1986), Glosten and Harris (1988), Menyah and Paudyal (2000) and Stoll (1989) relate the spread or the change in the spread to a vector of characteristics that are associated with the individual securities. These factors identified in prior spread models are the market value of the firm, the risk of the security, and the trading volume. On the other hand, the bid-ask spread also influences risk and trading activity in the financial market. Tobin (1978, 1984), Stiglitz (1989) and Summers and Summers (1989) argue that higher transaction cost leads to a lower market volatility since it discourages the short-term speculation which is believed as turmoil in the market. However, opposing views include Friedman (1953), Miller (1991), Dooley (1996) who believe that short-term speculation benefits to stabilizing as well. Grundfest and Shoven (1991) and Kupiec (1996) also point out that higher transaction cost leads to lower transaction frequency which requires larger price movement. A deep understanding of the two-way causality relationships between liquidity and risk certainly benefits to maintaining a healthy financial market.

The rest of this section outlines the structure of the thesis, and briefly summarizes the subsequent chapters and their main results. The following chapters represent a collection of five self-contained research articles and can be read independently.

Chapter 1 studies the impact of financial liberalization and political instability on economic growth. It contributes to this literature by using an innovative econometric methodology and unique data set: it presents power ARCH estimates for Argentina from 1896 to 2000. The main results show that the long-run effect of financial liberalization on economic growth is positive while the short-run effect is negative, albeit substantially smaller. The political instability effects are substantially larger in the short- than in the long-run. We also investigate potential mechanisms for the effects of financial liberalization and political instability on economic growth: direct impact or happening through the variation of growth volatility.

Chapter 2 quantitatively examines the relative importance of the identified underling reasons of Argentine riddle. Argentina is the only country in the world that was "developed" in 1900 and "developing" in 2000, various underlying reasons have been identified (chiefly political instability, financial development, inflation, trade openness, and international financial integration). We use the power-ARCH framework and annual data since 1896 to study how important are these factors vis-à-vis both growth and growth volatility. Our results suggest that financial development, trade openness and political instability are the main factors, with important differences in terms of their short versus long-run behavior.

Chapter 3 investigates real business cycle variability - growth relationship over a period which includes significant variation in output growth, such as the 1970s oil shock (supply shock), the 2001 electronic industry depression (demand shock) and the credit crunch in 2008 (financial crisis). Using monthly data in four Asian countries/regions (Japan, South Korea, Singapore and Taiwan) and parametric power ARCH methodology to proxy uncertainty, we obtain two important findings. First, we find that more uncertainty about output leads to a higher rate of growth in three of the four countries/regions and the form of the uncertainty is matter. Second, output growth reduces its uncertainty in all countries/regions via inflation uncertainty except Singapore. Our results are robust to alternative specifications and provide strong support to the recent emphasis by macroeconomists on the joint examination of economic growth and the variability of the business cycle.

Chapter 4 contributes to the issue of the welfare costs of inflation by examining the link between inflation and its uncertainty. We use parametric power ARCH models of the conditional variance of inflation to capture the relationship between inflation and its uncertainty by using monthly data for Japan, South Korea, Singapore and Taiwan over a period ranging from 1965 to 2010. For all countries/regions, inflation significantly raises inflation uncertainty as predicted by Friedman. Increased uncertainty affects inflation positively in Japan and Singapore in support of the Cukierman-Meltzer hypothesis. We find negative sign for Taiwan which is in accordance with the Holland hypothesis when error term is normal distributed, however the results are not statistically significant when student-t distribution is applied. Interestingly, South Korea's data reveals positive sign initially, however, it turns around when structural dummy is incorporated. This dramatic outcome in favour of Holland hypothesis, and chime in with Dueker and Kim (1999) who claim that the inflation was strictly controlled by South Korean monetary authority. In a sensitivity analysis we show that an arbitrary choice of the heteroskedasticity parameter influences this relationship significantly.

Chapter 5 investigates two-way causality relationships between spread, volatility and volume in FTSE100 stock index over the period from 1992 to 2004 by using bivariate AR-FI-GARCH model and multiple measurements of risk and spread. Measurements of spread include relative bid-ask spread, effective bid-ask spread, the inventory cost component of the bid-ask spread and the information cost component of the bid-ask spread. Risk is measured by two measurements of price volatility: the close-to-close volatility and the range-based volatility. We also take the impact of electronic trading into account. Our results suggest that spread and volume are positively impacted by volatility simultaneously. On another hand, both volatility and volume are negatively affected by spread. Furthermore, we find that the inventory cost component of spread has negative effect on volatility, in contrast, the information component of spread positively impacts on volatility. These results support the argument that speculations generate volatility in the market and higher transaction costs benefit stability of the market.

Chapter 1
Financial Liberalization, Volatility and the Finance-Growth Nexus

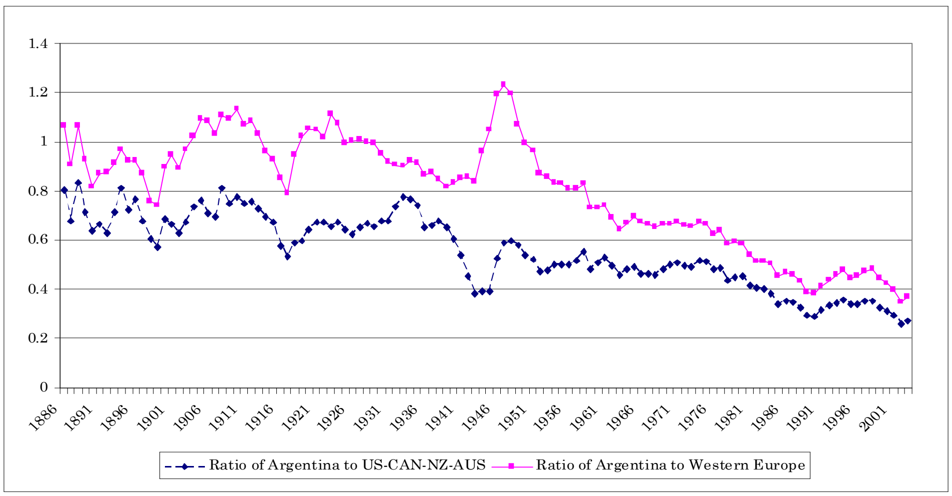
1.1 Introduction

Instability and performance are often inversely related. Financial crises are associated with growth decelerations and contractions, while political protest tends to disrupt productive activities thereby negatively affecting economic growth. Such amplified uncertainties, driven either by economic or political events, have deleterious consequences in terms of economic performance, especially in the short-run. In the long-run, however, financial development and political stability may instead have positive effects on growth. For example, the supply of credit to the private sector and transitions from autocracy to democracy are often considered key determinants of long-run growth across countries. In this light, this paper tries to answer the following questions. What is the relation between financial development on the one hand and economic growth and its volatility on the other? Do the sign and intensity of such effects vary over time and do they vary with respect to short- versus long-run considerations? Is there a dynamic asymmetry in the impact of financial development and political instability (that is, is it negative in the short- and positive in the long-run)?

This paper tries to tackle these questions using an innovative econometric framework and a unique type of data as it employs the power-ARCH (PARCH) framework and annual time series data for Argentina covering the period from 1896 to 2000. The "Argentinian puzzle," according to della Paolera and Taylor (2003), refers to the fact that since the Industrial Revolution, Argentina is the only country in the world that was developed in

1900 and developing in 2000 (see Figure 1.1)[1]. Campos et al. (2008) provide evidence that the development of financial markets as well as political instability are the most important factors in understanding this puzzle (these are more important than reasons normally given for the Argentinian puzzle such as inflation, trade openness and international financial integration.)

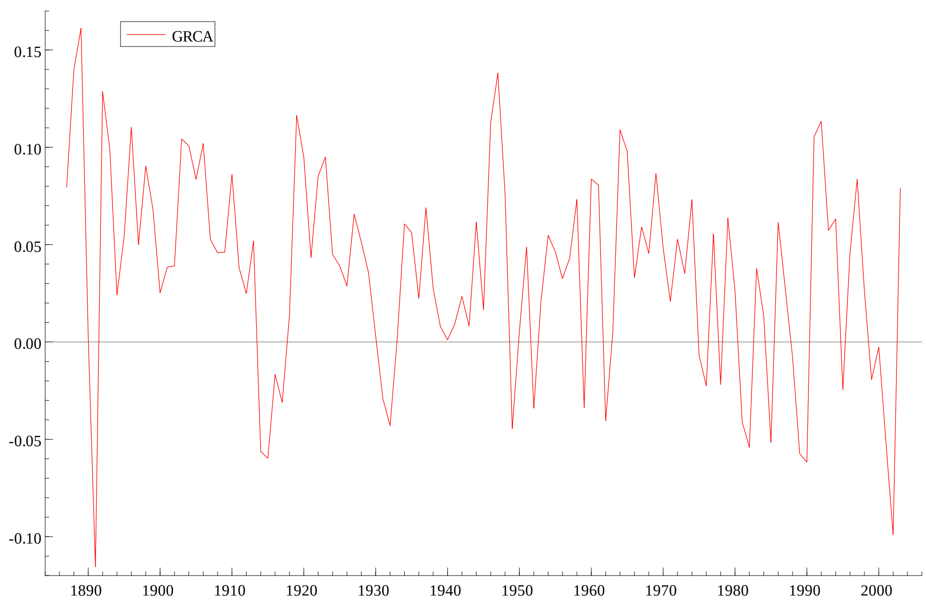
Fig. 1.1. Ratio of Argentina's GDP per Capita to Developed Countries' GDP per Capita, 1885-2003



As far as the literature on the finance-growth nexus is concerned, the present paper tries to contribute by offering econometric evidence based on historical data. Levine (2005) and Fecht et al. (2008) argue that the prevailing consensus favors a positive, lasting and significant effect from financial development to economic growth and that such effects are predictably stronger from measures of financial efficiency (for instance, the share in GDP of credit to the private sector) than from standard measures of financial depth (such as M3 over GDP). By using a range of financial development measures we can throw light on the impacts of these different dimensions over a period of time much longer than

[1] Authors' calculations using GDP per capita data from Maddison (2007), Western Europe is defined as Austria, Belgium, Denmark, Finland, France, Germany, Italy, Netherlands, Norway, Sweden, Switzerland and United Kingdom. The other group, Western Offshoots, includes Australia, Canada, New Zealand and United States.

Fig. 1.2. Real GDP Growth



that normally considered in the literature. Doing so, also allows us to investigate, inter alia, whether the impact of financial development on growth occurs directly or through growth volatility[2].

An important issue we tackle is that of the contrasting short- versus long-run effects of finance on growth. Seminal papers are those by Kaminsky and Schmukler (2003), Tornell et al. (2004) and Loayza and Rancière (2006). Despite the development of the financial system being robustly associated with economic growth, it is also often found to be the main predictor of financial crises. That is, while the long-run effect of finance on growth is positive, in the short-run it is negative. However, cross-country heterogeneity and business cycle synchronization issues may play an undesirably large role in generating this result and in particular regarding the relative magnitudes of these two effects. For instance, Loayza and Rancière (2006) report that the size of the effects is similar but

---

2 Levine (2005) surveys the finance and growth literature. On finance and volatility, see Bekaert et al. (2006) and Prasad et al. (2004).

the negative short-run effect is often larger than the positive long-run effect. In this paper we use data for a sufficiently long period of time and find supporting evidence for this asymmetric dynamic effect with the negative short-run effect being substantially smaller than the positive long-run effect[3]. Moreover, we try to shed light on important puzzles such as the one regarding the duration of the political instability effects. While the conventional wisdom is that these effects are severe in the long-run, Campos and Nugent (2002)[4] and Murdoch and Sandler (2004) argue that they are stronger in the short- than in the long-run.

One last intended contribution is to try to bridge the literature on the macroeconomics of instability (based on cross-sectional and short-panels) with that on the relationship between growth and its volatility, which is mostly time-series based[5]. The latter tends to downplay the potential dependence between growth and its volatility by assuming a linear relationship, the so called Bollerslev GARCH specification. Another final puzzle we try to address is on the sign of the growth-volatility link: while Grier and Tullock (1989) argue that larger standard deviations of growth rates are associated with larger mean growth rates, Ramey and Ramey (1995) show that output growth rates are adversely affected by their volatility.

Anticipating our main findings, we note the following in relation to the questions raised at the outset. The relationship between, on the one hand, financial development and political instability and economic growth, on the other, is not as straightforward as one may think at first. We find that it crucially depends on the type of political instability and of financial development, as well as short- versus long-run considerations. The short-run effect on economic growth of both informal instability and financial development is

---

[3] One important issue, which is beyond the scope of this paper, is regarding the causes of financial development, in particular, the legal origins versus political institutions debate (see Haber and Perotti, 2007).

[4] They argue that the long-run negative effect of political instability on growth depends on the inclusion of African countries and of institutions.

[5] Durlauf et al. (2005) survey the former, and Grier et al. (2004), Fountas and Karanasos (2006) and Fountas et al. (2007) review the latter. One paper that tries to link these literatures, and is close to ours in this sense, is Asteriou and Price (2001), which present time series (quarterly) data evidence for the UK since 1960.

negative and direct and these results are robust to accounting for structural breaks, which are important in light of the long time span we cover in this study. Yet, while the long-run influence of finance is positive, that of informal instability remains negative. We also find that the impact of formal instability is mostly indirect and operates through growth volatility. These results suggest that the "severity" of the political instability effects in a sense "dominates" that of financial development: while the short- and long-run finance effects work in opposite directions, the effects of political instability are both negative and seem to operate through different channels. In this paper, we show that formal political instability is detrimental to growth via the volatility channel and our results suggest that, together with informal instability, may have played a truly substantial role in the decline of the Argentinian economy during the XXth century.

The paper is organized as follows. Section 2 sets the context by showing how political instability and financial development contributed to the decline of Argentina from a position of a rich or developed country in year 1900 to that of a middle-income or developing country in year 2000. Section 3 describes the data. Section 4 details the econometric methodology. Section 5 discusses the main results. Section 6 concludes and suggests directions for future research.

1.2 The Role of Finance and Instability in Argentinian Growth

Among economic historians, there is little disagreement that the period from 1875 to the eve of World War I is the Belle Époque of Argentinian economic history (Taylor, 1992; Sanz-Villarroya, 2007). There is also little disagreement that Argentina's uniqueness derives from no other country having ever climbed so dramatically down from the selected group of developed countries (Figure 1). The two major disagreements remain not about whether but when the decline started and why. Some authors argue that it started with the 1930 crisis (e.g., Diaz-Alejandro 1985). Others argue for an earlier turning point (Taylor, 1992, argues for 1913), while Sanz-Villarroya (2005) argues for an even earlier structural break[6].

Irrespective of exactly when the decline started, until the immediate post II World War, Argentina was still ranked 10th country in the world in terms of per capita income (Alston and Gallo, 2007, p. 6). Della Paolera and Taylor (2003, p.5) note that "by 1900 Argentina's income per capita had risen from about 67 per cent of developed country-levels in 1870, to 90 percent in 1900, and 100 per cent in 1913. Whatever its exact status in 1913, for all practical purposes Argentina was an advanced country". They also calculate that after that the ratio of Argentina's income to OECD income fell to 84 percent in 1950, and then to 43 percent in 1987 (see also Figure 1.1). We calculate that this ratio rebounds in the 1990s but again reverts with the 2001 crisis[7]. It must not go unnoticed that in a recent book on the Great Depressions of the XXth Century (Kehoe and Prescott, 2007), Argentina is the only country that has two chapters (out of 16) entirely and solely dedicated to its economy.

It is not surprising, therefore, that there is a vast literature on the Argentine puzzle, providing alternative explanations for the long-run relative economic decline. Campos et al. (2008) provide a quantitative assessment of the relative importance of the causes that have been identified in the economic history literature (namely political instability or institutions, financial development, inflation rates, public deficits, trade openness and international financial integration). They find that the two most important ones are political instability and financial development.

A large number of studies underscore financial development as a major factor in the Argentinian puzzle (della Paolera and Taylor, 1998). Taylor (2003) associates the Argentinian decline with low savings rates (the high dependency rate linked to the liberal

---

6 Below we present and discuss our Bai-Perron estimates of the date of structural breaks in Argentinean growth. We find (and adjust our estimates accordingly) evidence for two structural breaks: 1922 and 1964.

7 Growth was negative from 1999 culminating with -10% in year 2002. The 2001 crisis entailed a default on large part of the external debt, devaluation, inflation, and the freezing of bank accounts (the corralito.) Riots, looting and anti-government demonstrations followed. See Kehoe (2003) for a discussion.

immigration policies). A related argument is that (restricted) access to finance perpetuates high levels of wealth and income inequality. More recently, Prados de la Escosura and Sanz-Villarroya (2009) have argued that the size and efficiency of financial intermediation ("contract intensive money") is key to explaining the Argentine puzzle.

Although a large literature associates the long-run relative decline of the Argentinian economy with political factors (see della Paolera and Taylor, 2003, and references therein), we are unaware of studies that try to evaluate this association quantitatively. For instance, Acemoglu and Robinson (2006, p.7) observe that: "The political history of Argentina reveals an extraordinary pattern where democracy was created in 1912, undermined in 1930, re-created in 1946, undermined in 1955, fully re-created in 1973, undermined in 1976, and finally reestablished in 1983". In a recent paper, Alston and Gallo (2007) identify the onset of widespread electoral fraud in the 1930s as a turning point for the erosion of the rule of law and thus as a major reason for the Argentinian decline.

In what follows, we take these considerations on board to provide a quantitative account of the relative importance of two of the main reasons often identified with the Argentinian debacle, political instability and financial development.

1.3 Measurement

Our data set contains measures of political instability, financial development and economic growth. The main data source is the Cross National Time Series Data set (Banks 2005), which has historical series on income per capita and various dimensions of instability. This is a commercial database that has been extensively used in the scholarship on growth, financial development and political instability (Durlauf et al., 2005.) Data are available yearly for Argentina from 1896 until 2000, excluding the years of the two World Wars.

Drawing upon the literature on growth and finance (Levine 2005) we use a broad range of measures of financial development, some reflecting depth and others efficiency aspects (see Figure 1.3). One note of caution is that there are various aspects of financial development which may be considered important but for which data are only available after about 1960 (e.g., intermediation spreads) and hence can not be used in the present study.

The first indicator we use is the ratio of M3 to GDP, from Alston and Gallo (2007). The main reason for considering this measure is that it has been used extensively in the finance-growth literature (King and Levine, 1993; Levine, 2005). One well-known drawback of this measure, however, is that the ratio of M3 to GDP reflects financial depth or the relative size of the financial system. It does not necessarily reflect how efficient the financial system actually is. We also use a narrower version of this variable (M1 over GDP) to check for the robustness of our results (source of the data is Bordo et al., 2001).

Our two other measures of financial development try to capture the efficiency of the financial sector, not its relative size. The source for both is Mitchell (2003). The first is the bank deposits by the private sector as a share of GDP (private deposits). A second measure is the total deposits in savings banks as a share of GDP. Given its more restrictive nature, we use the latter mostly for robustness checks, thereby attaching greater weight to private deposits[8].

We use a taxonomy of political instabilities based on the distinction between formal and informal (that is, whether or not instability originates from within the political system)[9]. Our informal political instability variables are as follows: annual number of anti-government demonstrations (peaceful public gatherings of at least 100 people), assassinations (defined as politically motivated murders or attempted murders of a high government official or politician), guerrilla warfare (armed activity, sabotage, or bombings by independent bands of citizens and aimed at regime overthrow), strikes (a general strike

[8] Because these financial development variables are found to be I(1), in the estimation they all enter in first-differences (see Figure 1.3).

[9] Our political instability variables enter one by one in the econometric estimation, thus they are not affected by the taxonomy itself. The taxonomy is introduced in Campos and Karanasos (2008).

Fig. 1.3. Measures of Financial Development

M3 over GDP

M1 over GDP

Private Deposits over GDP

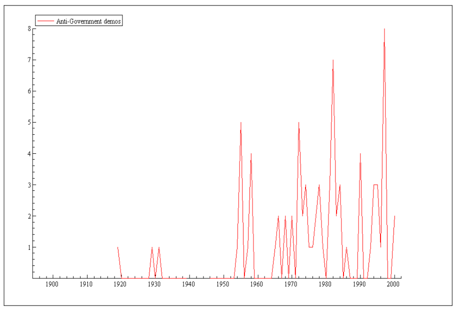
Savings Banks Deposits over GDP

of 1,000 or more workers involving more multiple employers and aimed at government policies), and revolutions (illegal or forced change in the top governmental elite, attempts at, or successful or unsuccessful armed rebellion). These series are available since 1919 (Figure 1.4). Our formal political instability variables (Figure 1.5) are as follows: the number of cabinet changes, the size of the cabinet, the number of constitutional changes, government crises, the number of legislative elections, and purges (which measure any systematic elimination by jailing or execution of political opposition within the ranks of the regime or the opposition)[10].

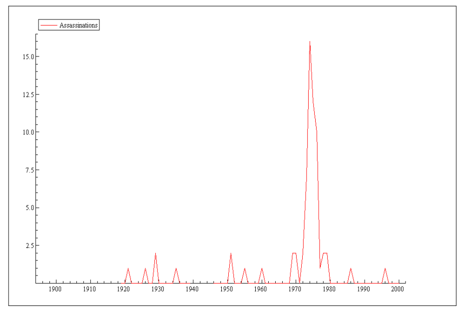
Before discussing methodological issues, we note that Granger causality and Hausman exogeneity tests were carried out and these support treating causality as flowing from financial development and political instability to economic growth, and not the other way around (these results are available upon request).

[10] Among all formal instability variables, "purges" is the closest to what we call informal instability, while "revolutions" is the one we think is closer to the formal instability variables.

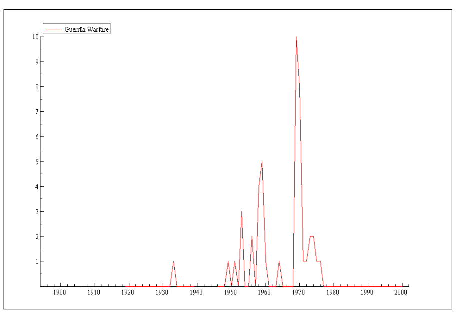
Fig. 1.4. Measures of Informal Political Instability



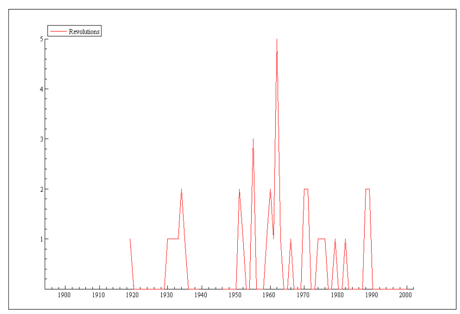
Anti-Government Demostrations



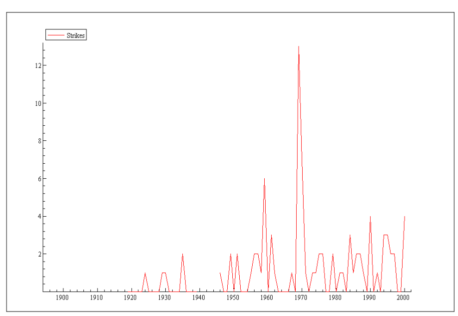
Assasinations



Guerilla Warfare

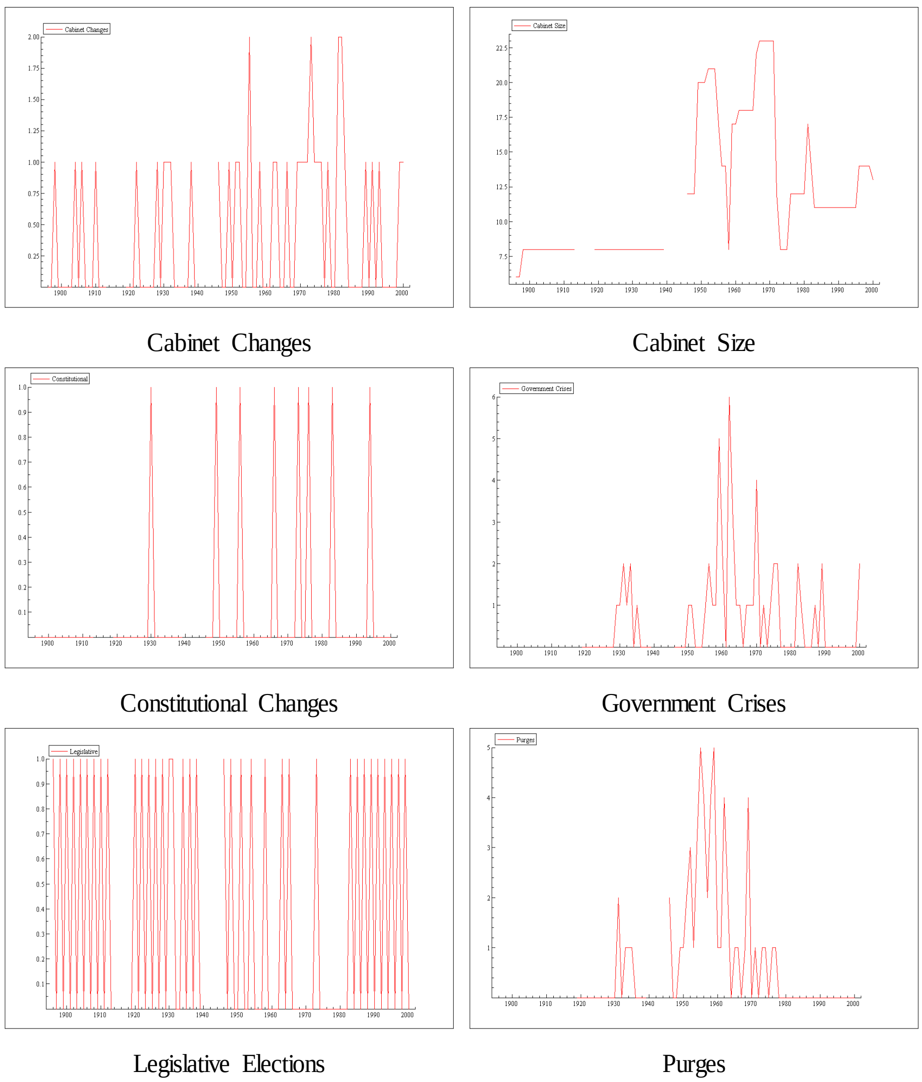


Revolutions



Strikes

Fig. 1.5. Measures of Formal Political Instability



1.4 Econometric Framework

The PARCH model was introduced by Ding, Granger and Engle (1993) and quickly gained currency in the finance literature[11]. Let growth (y_t) follow a white noise process augmented by a risk premium defined in terms of volatility (h_t):

$$y_t = c + kh_t + \lambda x_{it} + \epsilon_t, \tag{1.1}$$

with

$$\epsilon_t = e_t h_t^{\frac{1}{2}},$$

where x_{it} is either the political instability or the financial development variable.

In addition, $\{e_t\}$ are independently and identically distributed (i.i.d) random variables with $E(e_t) = E(e_t^2 - 1) = 0$, while h_t is positive with probability one and is a measurable function of the sigma-algebra $\sum_{t-1}$, which is generated by $\{y_{t-1}, y_{t-2}, \ldots\}$.

In other words, h_t denotes the conditional variance of growth. In particular, h_t is specified as an asymmetric PARCH(1,1) process with lagged growth included in the variance equation:

$$h_t^{\frac{\delta}{2}} = \omega + \alpha f\left(\boldsymbol{\varepsilon}_{t-1}\right) + \beta h_{t-1}^{\frac{\delta}{2}} + \gamma y_{t-l} + \phi x_{it}, \tag{1.2}$$

with

$$f\left(e_{t-1}\right) = \left[\left|e_{t-1}\right| - \varsigma e_{t-1}\right]^{\delta},$$

where δ (with $\delta > 0$) is the heteroscedasticity parameter, α and β are the ARCH and GARCH coefficients respectively, ς with $|\varsigma| < 1$ being the leverage term and γ being the level term for the lth lag of growth. In order to distinguish the general PARCH model

[11] See, for example, Karanasos and Kim (2006). Karanasos and Schurer, (2005, 2008) use this process to model output growth and inflation respectively.

from a version in which δ is fixed (but not necessarily equal to two) we refer to the latter as (P)ARCH.

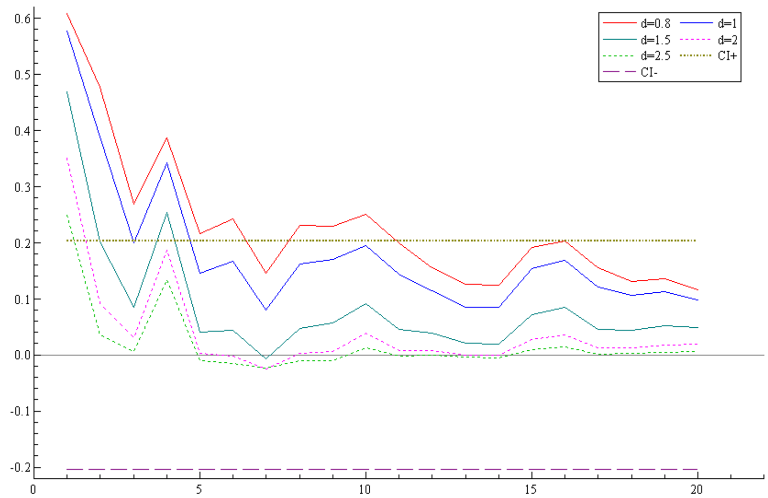
The PARCH model increases the flexibility of the conditional variance specification by allowing the data to determine the power of growth for which the predictable structure in the volatility pattern is the strongest[12]. This feature in the volatility process has important implications for the relationship between political instability, finance, inflation, and growth and its volatility. There is no strong reason for assuming that the conditional variance is a linear function of lagged squared errors. The common use of a squared term in this role is most likely to be a reflection of the normality assumption traditionally invoked. However, if we accept that growth data are very likely to have a non-normal error distribution, then the superiority of a squared term is unwarranted and other power transformations may be more appropriate.

The PARCH model may also be viewed as a standard GARCH model for observations that have been changed by a sign-preserving power transformation implied by a (modified) PARCH parameterization. He and Teräsvirta (1999) emphasize the point that if the standard Bollerslev type of model is augmented by the heteroscedasticity parameter (the power term), the estimates of the ARCH and GARCH coefficients almost certainly change.

Moreover, by squaring the growth rates one effectively imposes a structure on the data which may potentially yield sub-optimal modeling and forecasting performance relative to other power terms. One way to assess the severity of this assumption is to investigate the temporal properties of the power transformed absolute growth $|y_t|^d$. First, we examine the sample autocorrelations of the power transformed absolute growth $|y_t|^d$ for various positive values of d. Figure 1.6 shows the autocorrelogram of from lag 1 to 20 for a range of d values. The horizontal lines show the $\pm 1.96\sqrt{T}$ confidence interval (CI) for

[12] See Karanasos and Schurer (2005) and Karanasos and Kim (2006).

Fig. 1.6. Autocorrelation of $|y_t|^d$ from High to Low



the estimated sample autocorrelations if the process y_t is i.i.d. In this particular case, $CI = \pm 1.96\sqrt{T} = \pm 0.2032$.

The sample autocorrelations for $|y_t|^{0.8}$ are greater than the sample autocorrelations of $|y_t|^d$ for d= 1, 1.5, 2 and 2.5 at every lag. Or to put it differently, the most important conclusion from the autocorrelogram is that $|y_t|^d$ has the largest autocorrelation when $d = 0.8$. Furthermore, the power transformations of absolute growth when d is 0.8 have significant positive autocorrelations at least up to lag 10. Moreover, note that at all lags, $|y_t|^d$ has the lowest autocorrelation when d is 2 and 2.5. This result appears to argue against Bollerslev's specification.

Above all, the statistical significance of the in-mean effect is highly dependent on the choice of the value of the heteroscedasticity parameter. The effect might become insignificant if the power term surpasses a specific value. This suggests that if one assumes a priori a linear relationship between a variable and its uncertainty, the so-called Bollerslev specification, a significant link between the two might not be detected[13].

[13] Karanasos and Schurer (2008) find that the relationship between the variable and its conditional variance

1.5 Econometric Results

In this section we discuss our estimation results on the effects of political instability and financial development on economic growth in Argentina[14]. We start with a comparison between their direct and indirect effects on growth (the latter defined as taking place through growth volatility). We proceed by examining their asymmetric dynamic effects with emphasis on the economic significance of their long- and short-run impacts. This section closes with a discussion of a crucial robustness check (given the long span of time covered by our data), namely an assessment of the effects on our main results of accounting for structural breaks.

1.5.1 *Baseline Results*

We start with the estimation of the (P)ARCH(1,1) model in equations (1.1) and (1.2) in order to take into account the serial correlation observed in the levels and power transformations of our time series data. The tables below report the estimated parameters of interest for the period 1896-2000. These were obtained by quasi-maximum likelihood estimation (QMLE) as implemented in EVIEWS. The best fitting specification is chosen according to the Likelihood Ratio (LR) results and the minimum value of the Akaike Information Criteria (AIC) (not reported). Once heteroscedasticity in the conditional variance has been accounted for, our specifications appear to capture the serial correlation in the growth series[15].

---

is sensitive to changes in the values of the heteroscedasticity parameter. Put differently, the estimated values of the in-mean and the level effects are fragile to changes in the power term.

[14] As mentioned above, Campos et al. (2008) study the relative importance of the various factors often identified as main causes of the long-term relative decline of Argentinean per capita GDP since the late XIXth century. These include political instability (institutions), financial development, trade openness, international financial conditions, and inflation. They find that institutions (formal and informal political instability) and financial development are the most important factors.

[15] For all cases, we find that the leverage term is insignificant, so we re-estimate excluding this parameter. Controlling for both autoregressive terms as well as in-mean terms is important, because as shown in Ghysels et al. (2005), Conrad et al. (2010) and Conrad and Karanasos (2010) the omission of autoregressive terms may lead to spuriously significant in-mean terms. However, for all cases, we find that the IC choose the model

We specify model 1 with $\phi = \gamma = 0$ in equation (1.2) in order to study the direct effects of political instability and financial development, while model 2 with $\lambda = 0$ in equation (1.1) allows us to investigate their indirect impacts on growth. In all cases the estimates for the in-mean parameter (k) are statistically significant and positive. The estimated ARCH and GARCH parameters (α and β) are highly significant throughout[16].

For model 1 **(**$\phi = \gamma = 0$ **)**, when the informal political stability variables are used, the power term coefficient δ ranges from 0.8 (revolutions) to 1.0 (anti-government demonstrations). The corresponding value for all but one specification with formal instability variables is 0.8 (last column of Table 1.1). For model 2 (with $\lambda = 0$), with the informal instability variables AIC selects (P)ARCH models with δ equal to 0.9 (anti-government demonstrations, guerrilla warfare and strikes) or to 0.8 (assassinations and revolutions)[17]. For three out of the six formal variables the estimated value is 1 (last column of Table 1.2). Finally, for both models 1 and 2, when the financial development variables are used, in all but one case the IC chooses a (P)ARCH specification with estimated power term 0.8.

From the results for Model 1, the parameters λ for assassinations, guerrilla warfare and strikes (three measures of informal political instability) reveal their direct, negative and statistically significant impact on economic growth. Note also that none of the corresponding effects for the formal instability variables are statistically significant (Panel B). Importantly, we find the impact of financial development on economic growth to be positive and statistically significant, irrespective of the variable we use to measure it[18].

---

without the lagged growth and since its inclusion is insignificant, we re-estimate excluding this parameter. Moreover, the coefficients of lagged growth, which are always insignificant, do not qualitatively affect the main results below. These results are available from the authors upon request.

[16] With a limited number of observations the non-linear structure should not be overextended as this imposes excessive requirements on the data. Therefore we estimate the direct (model 1) and the indirect (model 2) effects separately.

[17] In the expressions for the conditional variances reported in Table 1.2, various lags of growth (from 1 to 12) were considered with the best model (l = 6) chosen on the basis of the minimum value of the AIC.

[18] We check the robustness of our findings with respect to the presence of level effects (results not reported). That is, we estimate model 1 with $\gamma \neq 0$. As with model 2 below, for all cases but one, there is evidence of a positive bidirectional feedback between growth and its volatility.

Table 1.1. Direct Effect of Political Instability and Financial Development on Economic Growth. (P)ARCH estimates

	k	λ	α	β	δ
Panel A. Informal Political Instability					
Anti-Government Demos	1.00 (1.66)	−0.0010 (1.36)	0.83 (4.15)	0.44 (1.90)	1.00 –
Assassinations	1.38 (1.85)	−0.0014 (1.70)	0.54 (3.86)	0.58 (4.04)	0.90 –
Guerilla Warfare	1.00 (3.69)	−0.0013 (4.35)	0.77 (5.43)	0.47 (3.13)	0.90 –
Revolutions	0.99 (4.31)	0.0001 (0.14)	0.61 (5.17)	0.59 (4.95)	0.80 –
Strikes	0.99 (3.07)	−0.0012 (2.13)	0.79 (4.66)	0.44 (2.38)	0.90 –
Panel B. Formal Political Instability					
Cabinet Changes	2.40 (3.97)	0.0001 (0.03)	0.31 (3.89)	0.72 (5.09)	0.80 –
Cabinet Size	0.79 (1.78)	0.0002 (1.47)	0.90 (4.13)	0.46 (2.32)	1.00 –
Constitutional	1.80 (1.99)	−0.0027 (1.35)	0.56 (3.01)	0.48 (1.25)	0.80 –
Government Crises	1.03 (2.53)	−0.0004 (0.42)	0.65 (4.89)	0.54 (3.59)	0.80 –
Legislative Elections	1.91 (2.69)	−0.0003 (0.15)	0.38 (3.43)	0.69 (5.79)	0.80 –
Purges	1.27 (1.92)	−0.0010 (1.55)	0.58 (4.03)	0.56 (3.69)	0.80 –
Panel C. Financial Development					
Private Deposits/GDP	0.76 (2.66)	0.98 (9.21)	0.70 (4.99)	0.57 (4.94)	0.80 –
Savings Bank Deposits/GDP	0.74 (1.80)	0.58 (3.43)	0.76 (4.36)	0.56 (4.97)	0.80 –
M3/GDP	0.81 (1.94)	0.32 (1.71)	0.94 (3.76)	0.43 (2.04)	1.00 –
M1/GDP	0.69 (2.30)	0.58 (5.14)	0.75 (5.29)	0.56 (5.85)	0.80 –

Note: This table reports parameter estimates for the following model:

$$y_t = c + kh_t + \lambda x_{it} + \varepsilon_t,\ h_t^{\frac{\delta}{2}} = \omega + \alpha \mid \varepsilon_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}}.$$

The numbers in parentheses are absolute t statistics.

The results in Table 1.1 are interesting for at least two reasons. One is that they provide evidence strongly suggesting that the type of political instability matters vis-à-vis economic growth: informal instability has a direct and negative effect, while formal instability does not. Second, they show that financial development has a positive and direct effect on growth, with M3 over GDP (a measure of the size of the financial sector) arguably being the weakest effect. In order to assess the robustness of these results, we investigated the effects of specifying instead lagged values of the informal instability measures and we have concluded that this does not qualitatively affect our main findings (these results are available from the authors upon request).

Examining the results for Model 2 (reported in Table 1.2) and focusing our attention on the ϕ and k parameters[19], we can now see that the formal political instability variables have strong indirect (through volatility) negative effects on growth. This result obtains for cabinet changes and size and constitutional changes. That is, these variables affect volatility negatively $(\phi < 0)$ and, since $k > 0$, they affect growth negatively as well[20]. Interestingly, none of the financial development and informal instability measures reveals such indirect effects (instead, as discussed above, they exhibit a direct impact on growth). These results reinforce the notion that the type of political instability matters with respect to economic growth: while informal instability may have a direct impact, the effect of formal political instability seems to operate indirectly, via growth volatility[21].

1.5.2 *Short-Run and Long-Run Effects*

In this section we investigate how short- and long-run considerations help us refine our baseline results. Another potential benefit is that the required use of lags may help allay any lingering concerns about endogeneity. In order to estimate short- and long-run relationships we employ the following error correction (P)ARCH form

$$\Delta y_t = \theta \Delta x_{i,t-l} + \varphi(y_{t-1} - c - \zeta x_{i,t-1}) + \varepsilon_t, \tag{1.3}$$

---

19 Note that, for all cases in model 2, there is evidence of a positive bidirectional feedback between growth and its volatility. The existing empirical literature focuses mainly on the effect of volatility on growth, see Fountas et al. (2006) and Fountas and Karanasos (2007).

20 We also estimate model 2 using an EGARCH specification. The results (not reported) are very much similar to the results we report in the paper.

21 Because data is normally missing for the War years both for economic growth and for the key explanatory variables (i.e., financial development and political instability), most of the available direct imputation methods would be ineffective, in that they would not be able to generate different results from the ones we report here. Moreover, we were able to obtain another, independent, GDP growths series from Aioli, Catao and Timmerman (2010) which differs from most other GDP series in that it contains information for the War years. The results (discussed below) we obtain from this other series are very much similar to the results we report in the paper.

Table 1.2. Indirect Effect of Political Instability and Financial Development on Economic Growth. (P)ARCH estimates

	k	α	β	γ	ϕ	δ
Panel A. Informal Political Instability						
Anti-Government Demos	1.25 (2.56)	0.65 (4.52)	0.46 (5.94)	0.17 (4.51)	−0.0002 (0.31)	0.90 –
Assassinations	1.09 (2.72)	0.68 (5.30)	0.27 (5.84)	0.27 (5.84)	−0.0038 (1.45)	0.80 –
Guerilla Warfare	1.12 (2.46)	0.73 (4.80)	0.46 (4.00)	0.10* (2.00)	0.0007 (0.82)	0.90 –
Revolutions	1.22 (2.03)	0.69 (3.73)	0.45 (2.37)	0.11* (1.80)	−0.0002 (0.14)	0.80 –
Strikes	1.14 (2.33)	0.70 (3.55)	0.48 (2.69)	0.06* (1.73)	0.0011 (1.27)	0.90 –
Panel B. Formal Political Instability						
Cabinet Changes	1.28 (1.96)	0.55 (2.99)	0.53 (4.90)	0.21 (3.93)	−0.0050 (4.03)	1.00 –
Cabinet Size	1.14 (1.77)	0.60 (3.89)	0.54 (5.02)	0.18 (2.48)	−0.0002 (2.13)	1.00 –
Constitutional Changes	1.18 (1.94)	0.69 (4.40)	0.45 (4.15)	0.18 (3.75)	−0.0077 (3.40)	1.00 –
Government Crises	1.12 (2.30)	0.72 (4.59)	0.47 (3.28)	0.11 (2.44)	0.0007 (0.57)	0.90 –
Legislative Elections	1.46 (2.52)	0.62 (4.70)	0.44 (3.52)	0.20 (3.12)	−0.0110 (1.17)	0.80 –
Purges	1.06 (3.00)	0.75 (5.26)	0.46 (4.32)	0.09 (1.99)	0.0004 (0.64)	0.90 –
Panel C. Financial Development						
Private Deposits/ GDP	2.05 (2.23)	0.41 (3.04)	0.62 (6.75)	0.40 (5.69)	0.58 (0.53)	0.80 –
Savings Bank Deposits/ GDP	1.81 (1.92)	0.53 (2.95)	0.60 (6.37)	0.32 (6.84)	−0.38 (0.38)	0.80 –
M3/GDP	1.35 (2.48)	0.58 (4.22)	0.48 (4.49)	0.23 (4.03)	0.29 (0.36)	0.80 –
M1/GDP	1.22 (2.56)	0.49 (5.30)	0.59 (9.46)	0.34 (5.67)	−0.13 (0.23)	0.80 –

Note: This table reports parameter estimates for the following model:

$$y_t = c + kh_t + \varepsilon_t,\ h_t^{\frac{\delta}{2}} = \omega + \alpha \mid \varepsilon_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}} + \gamma y_{t-6} + \phi x_{it}.$$

* * The orders of the lags are seven and five respectively.

The numbers in parentheses are absolute t statistics.

where θ and ζ capture the short and long-run effects respectively, and φ is the speed of adjustment to the long-run relationship.[22] This is accomplished by embedding a long-run growth regression into an ARDL model.[23] In other words, the term in parenthesis contains the long-run growth regression, which acts as a forcing equilibrium condition

$$y_t = c + \zeta x_{it} + u_t, \tag{1.4}$$

where u_t is $I(0)$. The lag of the first difference of either the political instability or financial development variables ($\Delta x_{i,t-l}$) characterizes the short-run effect. The condition for the existence of a long-run relationship (dynamic stability) requires that the coefficient on the error-correction term be negative and not lower than -2 (that is, $-2 < \varphi < 0$). We also take into account the (P)ARCH effects by specifying the error term ε_t as follows

$$\varepsilon_t = e_t h_t^{\frac{1}{2}}, \tag{1.5}$$

where

$$h_t^{\frac{\delta}{2}} = \omega + \alpha \left|\varepsilon_{t-1}\right|^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}}. \tag{1.6}$$

Table 1.3 presents the results on the estimation of short- and long-run parameters linking informal political instability or financial development with growth[24]. In all cases, the estimated coefficient on the error correction term (φ) lies within the dynamically stable range(-2,0). More precisely, the estimates of φ for informal instability and financial development lie within the range -0.71 to -0.50 and -0.85 to -0.44, respectively.

---

22 As pointed out by Loayaza and Rancière (2006) the requirements for the validity of this methodology are that: i) there exists a long-run relationship between the variables of interest and, ii) the dynamic specification of the model is sufficiently augmented so that the regressors are strictly exogenous and the resulting residual is serially uncorrelated.

23 For details on the “ARDL approach,” see Pesaran (1997) and Pesaran and Shin (1999).

24 In some cases where the routines did not converge we estimate the short- and long-run effects in two steps.

Regarding the short- and long-run effect estimates, θ and ζ, we focus our analysis first on those obtained from the informal instability variables. In all cases the estimates of the short-run coefficients are highly significant and negative and their absolute values are higher than the corresponding values for the long-run coefficients (for anti-government demonstrations, the long-run effect is not significantly different from zero). This provides evidence for the notion that the duration of the political instability effect does indeed matter and, for informal instability, such effects tend to be considerably stronger in the short- than in the long-run, in line with Campos and Nugent (2002) and Murdoch and Sandler (2004). The unexpected result is for revolutions: we found that the long-run effect on growth is positive. One possible explanation is that of escalation: political instability comes in cycles in which the level of political violence accelerates, with maxima coinciding with revolutions. Because revolutions reflect illegal or forced change in government elites (as well as successful or unsuccessful armed rebellions), their occurrence may be the culmination of a cycle of political violence, thus marking the beginning of a period of relatively low levels of political instability (and higher or more stable growth rates.) Another piece of evidence one can offer in support of this conjecture is that the revolutions series peaks around the date of the second structural break we identify in the GDP growth series (further details below.)

Next we discuss the important results regarding the financial development variables. In the long-run, we find that financial development affects growth positively ($\zeta > 0$). This result is in agreement with a large empirical literature (Levine 2005) and it is interesting that we reproduce it with a rather different methodology. However, the short-run coefficients tell a differently story: we find that the short-run impact of financial development on growth is negative and significant ($\theta < 0$). Thus our results square well with recent findings by Loayaza and Rancière (2006), among others, in that the sign of the relationship between economic growth and financial development crucially depends on the time horizon one takes (the effect being negative in the short- and positive in the longer-run.) It is also worth noting that our results are robust to various measures of financial development and also that the stronger long-run effects we obtain are for measures of financial efficiency rather than for measures of the size of the financial sector. Finally, as mentioned above, cross-country heterogeneity may play an undesirably large role in this type of result (e.g., Loayza and Rancière 2006) in that the size of the two effects this literature estimates is very similar, often with the negative short-run effect being somewhat larger. Because in this paper we use data for only one country and find supporting evidence for this asymmetric dynamic effect, one possible contribution is to help dispel such concerns about this result. We find that the negative short-run effect seems substantially smaller than the positive long-run effect (see also Table 1.6 below).

In sum, these results as a whole indicate that the severity of the political instability effects may "dominate" that of financial development: while the short- and long-run finance effects work in opposite directions, the effects of political instability are both negative and seem to operate through different channels. Formal political instability is detrimental to growth via the volatility channel and, together with informal instability (which has a negative direct effect on growth), may have played a truly substantial role in the relative decline of the Argentinian economy since its peak in the 1920s.

1.5.3 *Accounting for Structural Breaks*

In this section we subject our baseline results to an important robustness test. That is, we assess structural breaks. We use the methodology developed by Bai and Perron (2003) to examine whether there are any structural breaks in growth, its volatility, the various political instability and financial development variables[25]. This methodology is widely used because it addresses the problem of testing for multiple structural changes under very general conditions on the data and the errors. In addition to testing for the existence of breaks, these statistics identify the number and location of multiple breaks[26].

---

[25] Campos et al. (2009) compare a range of estimation methodologies for structural breaks in this type of low-frequency long historical data series.

[26] Details are available from the authors upon request.

Table 1.3. Short- and Long-run effects of Political Instability and Financial Development on Economic Growth. (P)ARCH estimates

	θ	ζ	φ	α	β	δ
Panel A. Informal Political Instability						
Antigovernment Demos	-0.0009 (2.92) $l=2$	0.0006 (1.04)	-0.70 (4.96)	0.98 (7.17)	0.41 (6.26)	0.80 –
Assassinations	-0.0019 (2.46) $l=4$	-0.0012 (3.52)	-0.71 (3.89)	0.82 (5.68)	0.53 (9.10)	0.90 –
Guerilla Warfare	-0.0014 (3.38) $l=3$	-0.0007 (2.59)	-0.60 (7.20)	1.10 (4.19)	0.36 (3.59)	0.90 –
Revolutions	-0.0015 (2.13) $l=5$	0.0013 (2.37)	-0.50 (3.60)	0.83 (5.76)	0.52 (6.85)	0.80 –
Strikes	-0.0026 (2.13) $l=4$	-0.0021 (2.74)	-0.54 (4.89)	0.76 (4.39)	0.55 (6.65)	0.80 –
Panel B. Financial Development						
Private Deposits/GDP	-1.35 (1.81) $l=5$	0.94 (23.72)	-0.44 (4.64)	0.37 (2.63)	0.80 (6.69)	0.90 –
Savings Bank Deposits/GDP	-0.55 (1.89) $l=1$	0.59 (4.84)	-0.70 (3.23)	0.74 (6.69)	0.56 (6.21)	0.80 –
M3/GDP	-0.16 (3.00) $l=4$	0.16 (1.60)	-0.83 (4.11)	0.81 (6.59)	0.52 (7.19)	0.80 –
M1/GDP	-0.21 (1.91) $l=1$	0.43 (4.20)	-0.85 (4.14)	0.74 (6.89)	0.54 (6.62)	0.80 –

Note: This table reports parameter estimates for the following model:

$\Delta y_t = \theta \Delta x_{i,t-l} + \varphi(y_{t-1} - c - \zeta x_{i,t-1}) + \varepsilon_t$, $h_t^{\frac{\delta}{2}} = \omega + \alpha \left|\varepsilon_{t-1}\right|^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}}$.

θ (l is the order of the lag) and ζ capture the short- and long-run effects respectively.

φ indicates the speed of adjustment to the long-run relationship.

The numbers in parentheses are absolute t statistics.

In the case of the economic growth series (and, interestingly, also for growth volatility) the Bai-Perron methodology supports two structural break points. The first occurs for year 1922 and the second for year 1964 (see figure 7). For our political instability variables, we find no structural breaks for the assassinations, guerilla warfare, cabinet and constitutional changes series[27], and nor do we find any breaks in the four financial development variables.

However, our Bai-Perron results support the idea that general strikes and government crises have one common structural break, which is dated at year 1955. This is the year of the military coup in which President Juan Domingo Perón was overthrown by the military. Breaks in the revolutions and purges series are detected for about the same political period, more specifically for year 1951 (see figure 7)[28]. Further, we also find one structural break in cabinet size and legislative elections (these are dated 1946 and 1949, respectively) while in anti-government demonstrations we find two breaks dated 1954 and 1972. With arguably one exception (anti-government demonstrations in 1972, which were motivated by demands for the return of Perón from exile), all the structural breaks in our political instability series occur during Perón governments. Perón was elected president three times. His first term is from 1946 to 1952. He is re-elected in 1951, his second term starts in 1952 and ends abruptly in 1955. His third term is between 1973 (where he is allowed to return from Spain after an18-year exile) and 1974 (when he suffers fatal heart attack.) Although marked by severe economic problems, the second term (1951 to 1955) is more often remembered for its political instability (the various terrorist attacks being a sad prelude to the so-called "Dirty War" of the 1970s).

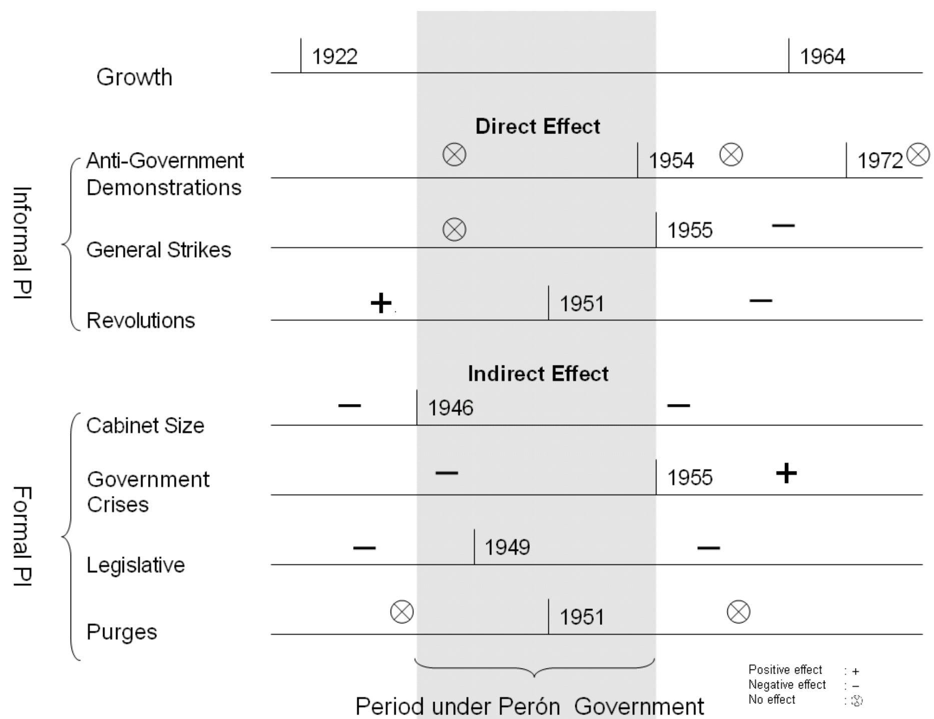
In what follows, we incorporate dummy variables in the equations (1.1) and (1.2), thus taking into account breaks in the political instability variables and in the volatility of

---

[27] Our data shows no guerilla warfare before 1948 and after 1977.

[28] In purges there is a second break dated 1978 but since after that year there were no purges we do not need to use a dummy variable to account for it.

Fig. 1.7. Structural Breaks



growth. First, we introduce the following notation. D_{1t}, D_{2t} are (intercept) dummies defined as D_{1t}, $D_{2t} = 1$ in the periods 1922-2000 and 1964-2000, respectively, and D_{1t}, $D_{2t} = 0$ otherwise. Similarly, D_{it} is a (slope) dummy indicating the period which starts from the year of the break in the political instability variable (x_{it}). For example for strikes and government crises $D_{it} = 1$ in the period from 1955 to 2000 whereas for cabinet size $D_{it} = 1$ during the period from 1946 until the end of the sample.
The augmented model is given by

$$y_t = c + kh_t + \lambda x_{it} + \lambda_1 D_{it} x_{it} + \epsilon_t, \tag{1.7}$$

and

$$h_t^{\frac{\delta}{2}} = \omega + \omega_1 D_{1t} + \omega_2 D_{2t} + \alpha \left|\varepsilon_{t-1}\right|^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}} + \gamma y_{t-l} + \phi x_{it} + \phi_1 D_{it} x_{it}. \tag{1.8}$$

Recall that the coefficients λ and ϕ capture the impacts of the political instability variable on growth and its volatility respectively. Similarly, λ_1 and ϕ_1 correspond to the two effects from the year of the break onwards. Thus the two effects are captured by λ and ϕ in the period up to the year of the structural break, and by $\phi + \phi_1$ and $\lambda + \lambda_1$ during the period from the year of the break until the end of the sample. As above in order to study the direct effects of political instability and financial development we specify model 1 with $\phi = \phi_1 = 0$, while model 2 with $\lambda = \lambda_1 = 0$ allows us to investigate their indirect impacts on growth.

We also incorporate intercept dummies and level effects in the conditional variance equation (1.6), as follows

$$h_t^{\frac{\delta}{2}} = \omega + \omega_1 D_{1t} + \omega_2 D_{2t} + \alpha \left|\varepsilon_{t-1}\right|^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}} + \gamma y_{t-l}. \tag{1.9}$$

Overall, we find our results to be robust to the inclusion of the structural break dummies (see Tables 1.4- 1.6)[29]. That is, (i) informal instability has a direct negative effect on growth, while formal instability has an indirect (through volatility) negative impact on growth, (ii) the effects of informal instability are significantly stronger in the short- than in the long-run, (iii) financial development affects growth positively in the long-run but negatively in the short-run, with the former being the dominant effect. As mentioned above, this latter result is very important. Previous research has robustly established that financial development affects growth positively in the long-run but negatively in the short-run but has struggled with the fact that the magnitude (that is, the economic significance) of the short-run effect tends to be larger than the long-run effect. This has been in part attributed to country heterogeneity. In this paper, we show that this may indeed be a major cause: the bottom panel of Table 1.6 shows that for all four measures of financial development we estimate the negative short-run effect to be smaller than the positive long-run effect. Further, for our preferred measure (private deposits), the magnitude of the long-run impact is more than three times that of the short-run effect (-0.29 versus 0.9).

Note also that the causal negative effect of strikes reflects the period 1955-2000 (see Panel A of Table 1.4), which is not surprising given the intricate relationship between the governments of Peron and organized labor. In addition, the impact of revolutions on growth becomes negative after 1951. As mentioned above, this is surprising and one possible explanation we offer is in terms of a cycle of escalation of political instability (which culminates in a revolution). It is also worth noting that before 1951, economic growth seems to be independent of changes in purges, whereas after 1951 a negative causal relationship starts, which implies that purges after 1951 behave similarly to the other informal instability variables (see Panel B of Table 1.4). Interestingly, the causal effects from legislative elections and government crises to growth volatility, become stronger after we account for their structural breaks in 1949 and 1955 respectively (see Panel B of Table

[29] These results are also robust to the inclusion of intercept dummies in the mean equation for growth (not reported).

1.5) with the latter result being the only unexpected one. Finally, note that when we take into account breaks and level effects in the volatility of growth, the long-run effects of assassinations and revolutions disappear (see Panel A of Table 1.6) thereby reinforcing our finding that the effects of informal political instability are more severe in the short- than in the long-run. Moreover, the coefficient of M3 over GDP also becomes insignificant, while the same does not happen to other, financial sector measures, in particular those reflecting efficiency (not size).

A final robustness test for our results was the use of alternative GDP growth series. Using the series constructed by Aiolfi et al. (2010) shown in Figure 1.2, we find that all our main conclusions regarding the direct and indirect effects remain unchanged. That is, (i) informal instability (anti-government demonstrations, guerilla warfare and strikes) have a direct negative effect on growth, while formal instability (constitutional, and legislative elections) have an indirect (through volatility) negative impact on growth, and (ii) financial development (private deposits/GDP and M3/GDP) has a positive and direct effect on growth. We also provide evidence for the notion that the duration of the political instability effect does indeed matter and, for informal instability, such effects tend to be significant only in the short-run but not in the long-run. Similarly, we find that the short-run impact of financial development on growth is negative and significant, whereas in the long-run the positive effect of financial development diminishes. Finally, we also note that we can not detect any structural break in the GDP growth series from Aiolfi et al. (2010) which provides additional support for our results in light of the above discussion.

Table 1.4. Direct Effect of Political Instability and Financial Development on Economic Growth: Accounting for Structural Breaks. (P)ARCH estimates

	k	λ	λ_1	ω_1	ω_2	α	β	γ	δ
Panel A. Informal Political Instability									
Anti-Government Demos	2.68 (2.00)	−0.003 (0.50)	0.003* (0.62)	−0.02 (1.44)	0.04 (3.40)	0.42 (3.20)	0.42 (3.21)	0.10 (2.90)	0.90 –
Assassinations	2.17 (1.85)	−0.002 (0.16)	–	–	0.04 (3.29)	0.48 (4.24)	0.37 (3.52)	0.09 (2.34)	0.90 –
Guerilla Warfare	2.07 (2.85)	−0.001 (2.68)	–	–	0.04 (3.54)	0.46 (5.34)	0.42 (5.40)	0.18 (4.68)	0.80 –
Revolutions	1.71 (2.63)	0.002 (4.19)	−0.003 (4.75)	−0.04 (1.82)	0.07 (3.52)	0.57 (5.40)	0.11 (1.12)	0.25 (5.70)	0.80 –
Strikes	2.17 (1.71)	0.0002 (0.38)	−0.001 (1.78)	–	0.03 (3.44)	0.49 (3.74)	0.41 (4.00)	0.06 (2.40)	1.00 –
Panel B. Formal Political Instability									
Cabinet Changes	2.41 (2.25)	0.0007 (1.05)	–	−0.02 (1.53)	0.06 (4.21)	0.41 (3.69)	0.36 (4.11)	0.10 (1.90)	0.80 –
Cabinet Size	2.12 (3.01)	−0.0006 (1.24)	0.0003 (0.89)	−0.02 (1.56)	0.05 (4.01)	0.44 (4.09)	0.42 (5.03)	0.16 (4.56)	0.80 –
Constitutional Changes	2.24 (2.13)	0.0007 (0.50)	–	−0.02 (1.70)	0.06 (4.21)	0.43 (3.28)	0.36 (2.74)	0.10 (1.99)	0.80 –
Government Crises	2.10 (1.70)	−0.0009 (1.14)	−0.0009 (1.08)	–	0.03 (2.85)	0.50 (3.86)	0.40 (3.95)	0.10 (4.70)	1.00 –
Legislative Elections	1.53 (1.69)	0.0010 (1.27)	−0.0030 (1.41)	–	0.02 (1.38)	0.50 (3.53)	0.47 (5.38)	0.22 (4.17)	1.00 –
Purges	2.43 (1.93)	0.0010 (1.59)	−0.0020 (2.87)	–	0.03 (3.26)	0.40 (3.23)	0.48 (6.20)	0.08 (6.51)	1.00 –
Panel C. Financial Development									
Private Deposits/GDP	2.26 (2.64)	0.90 (4.03)	–	−0.02 (1.58)	0.05 (4.33)	0.36 (2.72)	0.50 (4.04)	0.11 (1.79)	0.80 –
Savings Bank Deposits/GDP	3.75 (3.05)	0.63 (2.34)	–	–	0.03 (4.52)	0.51 (4.50)	0.45 (5.13)	0.04 (2.03)	1.00 –
M3/GDP	2.96 (1.89)	0.09 (0.35)	–	−0.01 (1.52)	0.04 (3.50)	0.38 (2.86)	0.43 (3.19)	0.04 (1.23)	1.00 –
M1/GDP	2.70 (2.15)	0.51 (2.85)	–	−0.02 (1.65)	0.05 (3.99)	0.39 (3.09)	0.43 (3.40)	0.06 (1.71)	0.90 –

Note: This table reports parameter estimates for the following model:

$y_t = c + kh_t + \lambda x_{it} + \lambda_1 D_{it} x_{it} + \varepsilon_t,$

$h_t^{\frac{\delta}{2}} = \omega + \omega_1 D_{1t} + \omega_2 D_{2t} + \alpha \mid \varepsilon_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}} + \gamma y_{t-6}.$

The numbers in parentheses are absolute t statistics.

*We include a second dummy (D_{it}) with estimated coefficient $0.0006(0.10)$.

Table 1.5. Indirect Effect of Political Instability and Financial Development on Economic Growth: Accounting for Structural Breaks. (P)ARCH estimates

	k	ω_1	ω_2	ϕ	ϕ_1	α	β	γ	δ
Panel A. Informal Political Instability									
Anti-Government Demos	2.06 (2.08)	–	0.04 (2.05)	0.006 (0.53)	−0.005* (0.43)	0.50 (3.32)	0.42 (3.06)	0.06 (1.21)	0.90 –
Assassinations	2.08 (3.22)	−0.03 (1.31)	0.05 (4.38)	−0.002 (1.11)	–	0.46 (4.74)	0.36 (3.22)	0.18 (4.07)	0.80 –
Guerilla Warfare	2.43 (1.71)	−0.03 (1.85)	0.05 (3.98)	0.001 (1.01)	–	0.43 (3.44)	0.39 (3.18)	0.08 (1.54)	0.90 –
Revolutions	1.93 (1.91)	−0.03 (2.34)	0.05 (4.03)	0.002 (2.57)	0.001 (0.62)	0.52 (4.49)	0.23 (2.83)	0.20 (6.49)	0.90 –
Strikes	2.38 (2.27)	–	0.04 (2.95)	0.001 (0.09)	0.001 (1.33)	0.41 (3.31)	0.47 (4.14)	0.06 (1.73)	0.90 –
Panel B. Formal Political Instability									
Cabinet Changes	2.29 (1.73)	−0.01 (1.51)	0.04 (4.03)	−0.003 (4.39)	–	0.43 (3.18)	0.33 (2.98)	0.12 (4.44)	1.00 –
Cabinet Size	2.73 (1.83)	−0.01 (1.66)	0.04 (4.25)	−0.001 (1.79)	0.0004 (1.58)	0.36 (3.00)	0.47 (5.06)	0.04 (1.56)	1.00 –
Constitutional Changes	2.84 (2.22)	−0.10 (1.42)	0.03 (3.57)	−0.020 (4.60)	–	0.30 (2.21)	0.53 (5.10)	0.07 (3.30)	1.00 –
Government Crises	2.75 (2.25)	–	0.04 (3.30)	−0.002 (1.80)	0.003 (3.50)	0.33 (2.66)	0.55 (4.74)	0.008 (0.03)	0.90 –
Legislative Elections	1.97 (2.23)	–	0.02 (1.21)	−0.010 (1.69)	−0.005 (2.43)	0.49 (3.66)	0.46 (3.14)	0.12 (2.67)	0.90 –
Purges	1.96 (1.38)	−0.04 (2.30)	0.06 (3.79)	0.003 (1.57)	−0.002 (1.11)	0.55 (5.05)	0.21 (2.37)	0.15 (5.02)	0.90 –
Panel C. Financial Development									
Private Deposits/GDP	2.89 (2.19)	−0.03 (1.71)	0.07 (4.07)	0.02 (0.02)	–	0.37 (2.80)	0.39 (2.16)	0.12 (1.89)	0.80 –
Savings Bank Deposits/GDP	2.37 (1.71)	–	0.05 (3.24)	−0.57 (0.61)	–	0.35 (2.43)	0.54 (3.37)	0.16 (3.23)	0.80 –
M3/GDP	1.78 (2.33)	–	0.04 (1.95)	0.59 (0.64)	–	0.62 (4.66)	0.29 (2.31)	0.24 (3.43)	0.80 –
M1/GDP	2.73 (2.73)	−0.02 (1.64)	0.06 (3.74)	−0.16 (0.17)	–	0.40 (3.16)	0.38 (2.53)	0.13 (2.87)	0.80 –

Note: This table reports parameter estimates for the following model:

$y_t = c + kh_t + \varepsilon_t$, $h_t^{\frac{\delta}{2}} = \omega + \omega_1 D_{1t} + \omega_2 D_{2t} + \phi x_{it} + \phi_1 D_{it} x_{it} + \alpha \mid \varepsilon_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}} + \gamma y_{t-6}$.

The numbers in parentheses are absolute t statistics.

*We include a second dummy (D_{it}) with estimated coefficient $-0.007(1.06)$.

Table 1.6. Short- and Long-run Effects of Informal Political Instability and Financial Development on Economic Growth: Accounting for Structural Breaks. (P)ARCH estimates

	θ	ζ	φ	ω_1	ω_2	α	β	γ	δ
Panel A. Informal Political Instability									
Antigovernment Demos	-0.0010 (1.92) $l=1$	-0.0006 (0.93)	-0.37 (3.62)	–	0.03 (1.68)	0.47 (2.05)	0.48 (2.70)	0.12 (3.97)	0.90 –
Assassinations	-0.0018 (1.90) $l=4$	0.0004 (0.54)	-0.31 (3.36)	–	0.04 (2.28)	0.52 (2.15)	0.39 (2.14)	-0.01 (0.27)	0.90 –
Guerilla Warfare	-0.0012 (1.61) $l=6$	-0.0008 (2.46)	-0.27 (3.11)	–	0.04 (2.69)	0.62 (3.20)	0.30 (2.31)	0.04 (0.83)	0.90 –
Revolutions	-0.0004 (1.77) $l=0$	-0.0002 (0.37)	-0.22 (2.10)	-0.03 (1.54)	0.06 (3.85)	0.57 (2.35)	0.28 (2.83)	-0.05 (1.25)	0.90 –
Strikes	-0.0012 (1.96) $l=6$	-0.0012 (2.35)	-0.28 (2.83)	–	0.06 (2.43)	0.62 (3.49)	0.30 (2.18)	0.04 (0.53)	0.80 –
Panel B. Financial Development									
Private Deposits/GDP	-0.29 (3.22) $l=5$	0.90 (3.50)	-0.26 (2.16)	-0.03 (1.96)	0.07 (4.87)	0.45 (1.79)	0.31 (2.81)	-0.02 (0.47)	0.90 –
Savings Bank Deposits/GDP	-0.54 (3.55) $l=5$	0.60 (2.57)	-0.26 (2.27)	–	0.05 (1.71)	0.58 (3.18)	0.41 (1.91)	0.06 (0.66)	0.80 –
M3/GDP	0.08 (0.58) $l=4$	0.11 (0.58)	-0.23 (2.26)	-0.02 (1.60)	0.04 (4.02)	0.55 (2.03)	0.30 (3.10)	-0.03 (1.00)	1.00 –
M1/GDP	-0.21 (2.78) $l=5$	0.35 (1.89)	-0.28 (2.73)	-0.03 (2.07)	0.07 (4.93)	0.45 (2.03)	0.34 (3.27)	-0.04 (0.69)	0.80 –

Note: This table reports parameter estimates for the following model:

$$\Delta y_t = \theta \Delta x_{i,t-l} + \varphi(y_{t-1} - c - \zeta x_{i,t-1}) + \varepsilon_t,$$

$$h_t^{\frac{\delta}{2}} = \omega + \omega_1 D_{1t} + \omega_2 D_{2t} + \alpha \left|\varepsilon_{t-1}\right|^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}} + \gamma y_{t-6}.$$

θ (l is the order of the lag) and ζ capture the short- and long-run effects respectively.

φ indicates the speed of adjustment to the long-run relationship.

The numbers in parentheses are absolute t statistics.

1.6 Conclusions

Within a power-ARCH framework using data for Argentina from 1896 to 2000, we find that: (a) informal political instability (assassinations, guerilla warfare, strikes) have a direct negative effect on economic growth, while formal instability (e.g., cabinet changes and size, and constitutional changes) have an indirect impact on growth (through its volatility); (b) financial development affects economic growth positively; (c) the informal instability effects are substantially larger in the short- than in the long-run; and (d) the financial development effects are negative in the short- but positive and substantially larger in the long-run.

These findings raise a number of new questions that we believe may be useful in motivating future research. Here we highlight two related suggestions: one on the role of finance and one on methodology. Regarding the role of finance in the process of economic development, our findings extend a large body of previous research in that we also show a strong, positive impact of financial development on growth in the long-run. However, the negative effects of political instability on growth might outweigh the positive influence of financial development. We find that different forms of political instability affect growth through different channels over different time windows, making up for a strong and rather resilient effect that seems very powerful vis-à-vis the benefits from financial development. Yet Argentina is unique: no other country in the world since the Industrial Revolution went from riches to rags. Put differently, Argentina is an outlier and further research should replicate our analysis using the historical experience of other countries (ideally in a panel setting). That is, future studies should focus on the relationship between political instability, financial development and economic growth in a panel of developing countries. Notice, however, that the data requirements are very heavy indeed, with most developing countries lacking historical data even on key figures, such as per capita GDP, going back to the beginning or middle of the XIXth century. This, of course, does not make this task less important. The second suggestion refers to a possible methodological improvement, namely the application of the bivariate PARCH model to the problem at hand (despite the relatively small number of observations). A joint estimation of the political instability-financial development-growth system in a panel of countries would clearly represent progress and is something we feel future research should try to address.

1.A Appendix

Table 1.7. Direct Effect of Political Instability/Financial Development on Economic Growth, with Laged Growth in Variance Equation. (P)ARCH estimates

	k	λ	α	β	γ	δ
Panel A. Informal Political Instability						
Anti-Government Demos	1.27 (2.55)	−0.0009 (1.47)	0.66 (4.66)	0.42 (4.47)	0.25 (5.61)	0.80 –
Assassinations	1.18 (2.60)	−0.0009 (1.70)	0.72 (4.63)	0.42 (4.03)	0.15 (2.80)	0.90 –
Guerilla Warfare	1.04 (2.38)	−0.0011 (2.88)	0.74 (4.11)	0.47 (5.26)	0.10 (6.08)	1.00 –
Revolutions	0.99 (1.98)	−0.0006 (2.22)	0.63 (6.38)	0.48 (7.06)	0.31 (7.54)	0.80 –
Strikes	1.22 (2.69)	−0.0012 (2.33)	0.66 (4.04)	0.43 (4.48)	0.23 (4.46)	0.90 –
Panel B. Formal Political Instability						
Cabinet Changes	1.23 (1.92)	0.0014 (1.32)	0.59 (3.94)	0.52 (5.06)	0.24 (4.09)	0.90 –
Cabinet Size	1.29 (1.76)	−0.0001 (0.29)	0.53 (2.97)	0.57 (5.34)	0.16 (2.85)	1.00 –
Constitutional Changes	1.28 (2.04)	−0.0031 (1.46)	0.54 (3.86)	0.54 (6.21)	0.23 (4.15)	0.90 –
Government Crises	1.24 (2.74)	−0.0030 (0.44)	0.62 (4.40)	0.47 (6.82)	0.20 (4.57)	0.90 –
Legislative Elections	1.21 (2.09)	0.0013 (1.30)	0.57 (4.02)	0.53 (8.72)	0.22 (5.82)	1.00 –
Purges	1.23 (2.12)	−0.0007 (1.21)	0.68 (3.62)	0.45 (4.82)	0.13 (3.95)	1.00 –
Panel C. Financial Development						
Private Deposits/GDP	0.96 (2.83)	0.84 (3.65)	0.53 (3.07)	0.61 (4.64)	0.31 (2.76)	0.80 –
Savings Bank Deposits/GDP	0.96 (2.14)	0.52 (3.43)	0.74 (3.91)	0.50 (4.17)	0.22 (2.89)	0.80 –
M3/GDP	1.06 (1.91)	0.47 (2.70)	0.65 (2.97)	0.51 (2.98)	0.05 (0.61)	0.80 –
M1/GDP	0.88 (3.04)	0.50 (3.52)	0.61 (3.99)	0.58 (5.99)	0.21 (3.86)	0.90 –

Note: This table reports parameter estimates for the following model:

$$y_t = c + kh_t + \lambda x_{it} + \varepsilon_t,\ h_t^{\frac{\delta}{2}} = \omega + \alpha \mid \varepsilon_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}} + \gamma y_{t-6}.$$

The numbers in parentheses are absolute t statistics.

Comment: The coefficient of the level effect is significant in all cases (see the γ column). The impact of the informal instability and financial development on growth is robust to the presence of level effects (see the λ column).

Table 1.8. Direct Effect of Informal Political Instability/Financial Development on Economic Growth, with Laged Growth in Mean and Variance Equations. (P)ARCH estimates

	k	ρ	λ	α	β	γ	δ
Panel A. Informal Political Instability							
Anti-Government Demos	0.27 (0.30)	0.29 (0.89)	−0.0009 (1.29)	0.58 (3.93)	0.54 (7.25)	0.19 (3.23)	0.90 –
Assassinations	0.45 (1.80)	0.27 (0.63)	−0.0009 (1.88)	0.58 (3.12)	0.50 (6.03)	0.26 (4.29)	0.80 –
Guerilla Warfare	1.05 (1.11)	0.05 (0.25)	−0.0011 (2.38)	0.57 (3.31)	0.55 (6.60)	0.12 (9.18)	1.00 –
Revolutions	0.73 (1.37)	0.10 (0.69)	−0.0005 (1.16)	0.60 (4.25)	0.51 (10.23)	0.20 (5.85)	0.90 –
Strikes	0.48 (2.21)	0.19 (1.36)	−0.0013 (3.44)	0.83 (5.30)	0.37 (5.49)	0.22 (4.04)	0.80 –
Panel B. Financial Development							
Private Deposits/GDP	0.79 (1.07)	0.09 (0.43)	0.71 (3.55)	0.45 (3.05)	0.67 (6.28)	0.33 (4.17)	0.80 –
Savings Bank Deposits/GDP	0.82 (0.83)	0.17 (0.97)	0.87 (3.47)	0.60 (.3.16)	0.59 (6.84)	0.30 (7.78)	0.80 –
M3/GDP	1.17 (1.45)	0.13 (0.73)	0.12 (0.79)	0.41 (2.46)	0.63 (4.89)	0.27 (4.18)	0.90 –
M1/GDP	0.96 (1.85)	0.05 (0.37)	0.44 (2.51)	0.47 (3.64)	0.63 (7.14)	0.34 (3.94)	0.80 –

Note: This table reports parameter estimates for the following model:

$$y_t = c + kh_t + \rho y_{t-1} + \lambda x_{it} + \varepsilon_t, h_t^{\frac{\delta}{2}} = \omega + \alpha \mid \varepsilon_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}} + \gamma y_{t-6}.$$

The numbers in parentheses are absolute t statistics.

Comment: In all cases the coefficient of lagged growth is insignificant (see the ρ column). Informal instability (assassinations, guerilla warfare and strikes) has a negative effect on growth whereas financial development (with the exception of M3/GDP) affects growth positively (see the λ column).

Table 1.9. Indirect Effect of Formal Political Instability on Economic Growth, with Laged Growth in Mean and Variance Equations. (P)ARCH estimates

	k	ρ	α	β	γ	ϕ	δ
Formal Political Instability							
Cabinet Changes	0.62 (1.30)	0.14 (0.78)	0.60 (2.73)	0.51 (3.37)	0.33 (3.33)	−0.0068 (1.79)	0.80 –
Cabinet Size	0.91 (1.94)	0.03 (0.17)	0.66 (4.30)	0.50 (5.17)	0.28 (3.73)	−0.0002 (1.16)	0.90 –
Constitutional Changes	0.81 (1.35)	0.17 (1.06)	0.42 (2.62)	0.63 (5.16)	0.24 (2.93)	−0.0168 (2.81)	0.80 –
Legislative Elections	0.31 (0.56)	0.39 (0.53)	0.51 (3.40)	0.52 (6.45)	0.16 (4.13)	−0.0193 (2.73)	0.90 –

Note: This table reports parameter estimates for the following model:

$$y_t = c + kh_t + \rho y_{t-1} + \varepsilon_t,\ h_t^{\frac{\delta}{2}} = \omega + \alpha \mid \varepsilon_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}} + \gamma y_{t-6} + \phi x_{it}.$$

No convergence for Government crises and Purges.

The numbers in parentheses are absolute t statistics.

Comment: In all cases the coefficient of lagged growth is insignificant (see the ρ column). Formal instability (cabinet and constitutional changes and legislative elections) has a negative effect on the volatility of growth (see the ϕ column).

Table 1.10. Indirect Effect of Formal Political Instability on Economic Growth, EGARCH estimates

	k	α	β	γ	ϕ
Formal Political Instability					
Cabinet Changes	$0.14^{\dagger}$ (1.07)	1.27 (5.74)	0.82 (19.52)	4.80 (2.93)	-0.85 (2.16) $l=2$
Cabinet Size	2.04 (3.19)	1.02 (7.40)	0.87 (17.73)	3.99 (3.80)	-0.004 (0.15) $l=2$
Constitutional Changes	1.48 (1.83)	0.94 (7.76)	0.90 (27.14)	2.64 (1.84)	-1.83 (2.56) $l=0$
Government Crises	0.49 (2.93)	2.60 (7.42)	0.79 (21.85)	6.61 (12.67)	-0.70 (9.17) $l=1$
Legislative Elections	2.69 (3.42)	1.30 (3.49)	0.78 (15.59)	4.41 (4.27)	-1.93 (5.55) $l=3$
Purges	$0.01^{\ddagger}$ (1.08)	1.27 (5.74)	0.82 (19.52)	4.80 (2.93)	-0.85 (2.16) $l=2$

Note: This table reports parameter estimates for the following model:

$y_t = c + kh_t + \varepsilon_t,$

$log(h_t) = \omega + \alpha \mid e_{t-1} \mid + \beta log(h_{t-1}) + \gamma y_{t-6} + \phi x_{it-l}.$

†Estimates for standard deviation.

‡Estimates for logarithm of variance.

The numbers in parentheses are absolute t statistics.

Comment: In all but one cases formal instability has a negative effect on the volatility of growth (see the ϕ column).

Table 1.11. Direct and Indirect Effects of Political Instability/Financial Development on Economic Growth (Real GDP Growth). (P)ARCH estimates

Panel A. Direct Effect: $\phi = \phi_1 = 0$	k	ρ	α	λ	λ_1	δ
Informal Political Instability						
Anti-Government Demos	5.73 (1.43)	0.23 (2.67)	0.42 (4.15)	−0.018 (7.49) $l=4$	0.017* (4.78) $l=4$	1.00 –
Guerilla Warfare	8.22 (0.67)	0.27 (2.76)	0.18 (1.75)	−0.005 (2.07) $l=3$	–	1.10 –
Strikes	12.09 (0.89)	0.27 (2.77)	0.18 (1.80)	−0.006 (3.77) $l=3$	–	1.00 –
Financial Development						
Private Deposits/GDP	15.79 (1.13)	0.29 (2.95)	0.19 (1.90)	0.07 (2.22) $l=1$	–	1.10 –
M3/GDP	7.15 (0.60)	0.31 (3.38)	0.18 (1.58)	0.18 (2.19) $l=3$	–	1.10 –
Panel B. Indirect Effect: $\lambda = \lambda_1 = 0$	k	ρ	α	ϕ	ϕ_1	δ
Formal Political Instability						
Cabinet Changes	0.94 (0.18)	0.24 (2.39)	0.25 (2.52)	−0.008 (2.56) $l=5$	–	1.10 –
Constitutional Changes	0.65† (1.57)	0.27 (2.91)	0.30 (2.76)	−0.011 (2.91) $l=4$	–	1.20 –
Legislative Elections	2.03† (1.75)	0.32 (4.10)	0.28 (2.15)	−0.006 (1.67) $l=2$	−0.004 (0.99) $l=2$	1.00 –

Note: This table reports parameter estimates for the following model:

$y_t = c + kh_t + \rho y_{t-1} + \lambda x_{i,t-l} + \lambda_1 D_{it-l} x_{it-l} + \varepsilon_t,$

$h_t^{\frac{\delta}{2}} = \omega + \alpha \mid \varepsilon_{t-1} \mid^{\delta} + \phi x_{it-l} + \phi_1 D_{it-l} x_{it-l}.$

*We include a second dummy (D_{it}) with estimated coefficient $-0.002(0.53)$.

†estimates for standard deviation.

The numbers in parentheses are absolute t statistics.

Comment: Informal instability (anti-government demos, guerilla warfare and strikes) has a negative effect on growth whereas financial development (private deposits/GDP and M3/GDP) affects growth positively (see the λ column in panel A). Formal instability (cabinet and constitutional changes and legislative elections) has a negative effect on the volatility of growth (see the ϕ column in panel B).

Table 1.12. The short- and long-run effects of Informal Political Instability/Financial Development on Economic Growth (Real GDP Growth)

	θ	ζ	ϕ
Panel A. Informal Political Instability			
Anti government Demos	-0.002 (0.43) $l=4$	-0.012 (0.26)	-0.71 (5.30)
Guerilla Warfare	-0.007 (2.86) $l=3$	-0.007 (2.06)	-0.68 (6.02)
Strikes	-0.004 (1.70) $l=3$	-0.006 1.71	-0.67 (6.18)
Panel B. Financial Development			
Private Deposits/GDP	-0.02 (0.83) $l=3$	-0.05 (1.26)	-0.69 (6.19)
M3/GDP	-0.10 (1.82) $l=5$	0.06 (0.54)	-0.69 (6.10)

Note: This table reports parameter estimates for the following model:

$$\Delta y_t = \theta \Delta x_{i,t-l} + \phi(y_{t-1} - c - \zeta x_{i,t-1}) + \varepsilon_t,$$

θ (l is the order of the lag) and ζ capture the short- and long-run effects respectively. ϕ indicates the speed of adjustment to the long-run relationship. The numbers in parentheses are absolute t statistics.

Comment: The short- and long-effects of informal instability (guerilla warfare and strikes) on growth are negative (see the θ and ς columns in Panel A). Financial development (M3/GDP) affects growth (negatively) only in the short-run.

Chapter 2
Non-Linear Econometric Evidence on Growth and Volatility in Argentina since 1890

2.1 Introduction

The general economic trend since the Industrial Revolution has clearly been one of economic betterment. Since 1850, a sustained increase in living standards is evident across the globe. Comparing the situation in 1900 with that in 2000, one can identify four different types of country trajectories. A handful of countries were rich or developed in, say, year 1900, and remain rich or developed in year 2000 (for example, the U.S. and the U.K.) A few other countries were developing in 1900, but turned around and by year 2000 were among the developed countries. Examples of this second group are Japan and most of the European periphery (including Portugal, Italy and Spain.) The vast majority of countries belongs to a third group of countries, those that were relatively poor in year 1900 and remain relatively poor or developing in year 2000. The fourth group of countries encompasses those that were developed in 1900 and are developing in 2000. Only one country falls into this category and that is Argentina.

Although placed among the highest incomes per capita in the world in 1900, "Argentina's ratio to OECD income fell to 84 percent in 1950, 65 percent in 1973, and a mere 43 percent in 1987…Argentina is therefore unique (della Paolera and Taylor, 2003, p. 5). Unsurprisingly, this "Argentine puzzle" has received a great deal of attention and scholars have identified several potential reasons, chiefly among them financial development, political instability (or institutions), macroeconomic volatility, inflation, trade openness, public deficit, and international financial integration. Surprisingly, we find no studies trying to quantify and assess the relative importance of this array of reasons. Hence this paper tries to fill this gap.

Within a power-ARCH (PARCH) framework and using annual time series data for Argentina covering the period from 1896 to 2000, the aim of this paper is to put forward answers to the following questions. What is the relationship between, on the one hand, financial development (domestic and international), public deficits and inflation, trade openness and political instability and, on the other, economic growth and volatility? Are the effects of these variables direct (on economic growth) or indirect (via the conditional growth volatility)? Does the intensity and sign of these impacts vary over time? Does the intensity of these effects vary with respect to short- versus long-run considerations? Is the intensity of these effects constant across the different eras or phases of Argentine economic history (in other words, are they independent from the main structural breaks we estimate)?

This paper tries to contribute to our understanding of the main causes of economic growth. Durlauf et al. (2005) and Acemoglu (2009) provide recent, authoritative surveys that support the view that there seems to be dissatisfaction with the empirical growth literature. This paper tries to improve matters in this regard by focusing on one the most undisputed and intriguing outliers (as opposed to follow the common practice of trying to learn something about growth by focusing on the mean or median country). We believe this study can further our understanding about economic growth because: (a) we study only one individual country over a very long period of time with annual frequency data, (b) we extensively use the economic history literature to guide our choice of potential important reasons for the Argentine decline, (c) we choose an econometric methodology that has been seldom used in the empirical growth literature despite the fact that it easily allow us to contrast the direct to the indirect (i.e., via the volatility channel) effects of each of our candidate reasons, sort out the short- from the long-run impacts, and distill the consequences of accounting for important structural breaks on the robustness of our key results.

Another important benefit of our choice of econometric framework is that it helps shedding light on an important and resilient puzzle on the relationship between output growth and its volatility. While Ramey and Ramey (1995) show that growth rates are adversely affected by volatility, Grier and Tullock (1989) argue that larger standard deviations of growth rates are associated with larger mean rates. The majority of ARCH papers examining the growth-volatility link are restricted to these two key variables. That is, they seldom assess whether the effects of the presence of other variables affect the relation and, in the rare occasions that happens, they are usually inflation and its volatility that comes into play.[30] One contribution of this paper is to study if and how the growth-volatility relationship changes in light of a much wider set of variables. Note also that the use of annual data allows us to perform a more appropriate test of the hypothesis that predicts a positive effect of output variability and uncertainty on the growth rate of output.[31]

Our results are presented following specific types of effects. That is, we discuss direct (on mean economic growth), indirect (via volatility), dynamic (short and long-run) and structural break effects. Moreover, in trying to satisfy both the time-series and economic growth literature traditions (the former mostly univariate and the latter multivariate), for each effect we report estimates for one variable at a time before discussing the full multivariate results.

As for the individual univariate direct effects on economic growth, we find evidence for direct influences on real GDP growth from financial development (ratio of private de-

---

[30] For a comprehensive review of this literature see Fountas et al. (2006). In addition, Gillman and Kejak (2005) bring together for comparison several main approaches to modeling the inflation-growth effect by nesting them within a general monetary endogenous growth model with both human and physical capital.

[31] Black (1987) argues that investments in riskier technologies will be pursued only if the expected return on these investments (expressed as the average rate of output growth) is large enough to compensate for the extra risk. As real investment takes time to materialize, such an effect would be more likely to obtain in empirical studies utilizing low-frequency data.

posits to GDP), informal political instability,[32] international financial markets (interest rate in the United Kingdom), trade openness and public deficit. Equally importantly, we find no such evidence from formal political instability and inflation rates. The multivariate analysis helps to narrow down this set to the positive effect of financial development (private and savings banks deposits to GDP) and to the negative effect of informal political instability (guerilla warfare and general strikes), international financial integration and trade openness as major drivers of growth in Argentina since the 1890s.

How does this set of variables affect predicted growth volatility? Or in other words, how do they affect growth indirectly through their impact on growth volatility)? The strongest indirect impacts we find from the univariate evidence are the volatility-increasing effects of inflation, public deficit and the UK interest rate and the volatility-decreasing effects of trade openness and constitutional changes. Importantly, however, our multivariate results show that the most robust indirect effects are those from the last three variables.

Our investigation of the dynamic effects shows important differences in terms of short and long run behavior of our key variables: informal political instability, public deficit and the UK interest rate affect growth negatively both in the short- and the long-run. The effects of the former are larger in the short- than in the long-run. Those for financial development and trade openness are negative in the short- but positive in the long-run. However, the effets of public deficit and trade openness disappear in the multivariate analysis.

Finally, we subjected all these results to the presence of structural breaks. This is a crucial exercise given the very long-term nature of our data. We find that our basic results remain once we take into structural breaks into account, the notable exception being that the (in)direct effects of (public deficit)UK interest rates disappear once breaks are accounted for.

[32] We follow Campos and Karanasos (2008) in separating informal from formal poltical instability. The former is defined as instability outside of the government realm and measured here as guerilla warfare, while the latter is defined as instability within the government realm and measured in this paper by the number of constitutional changes.

In sum, our main results suggest that financial development and political instability (or institutions) exhibit the most robust first-order effects on growth and volatility. We also find that trade openness and international financial integration (proxied in this paper by the UK interest rate) play important yet secondary roles because the effects of the former do not extend to the long-run (that is, they are restricted to the short-run) and those of the latter (direct effects) vanishing when structural breaks are fully accounted for.

The paper is organized as follows. Section 2 sets the historical context for the paper by documenting the decline of Argentina from a position of a rich or developed country in year 1900 to that of a middle-income or developing country in year 2000. More importantly, this section briefly reviews the Argentinean historiography stressing the main reasons that have been offered to explain the relative decline. Section 3 describes the data and Section 4 provides details and justification for our econometric methodology. Section 5 has our baseline econometric results and the results from various sensitivity tests are presented in Section 6. Section 7 concludes and suggests directions for future research.

2.2 The Argentine Riddle

Argentina was part of the Spanish colonial empire for about three centuries. Its name as well as the name of its main river indicates that the colonizers had clear expectations: they expected it to become one of the main transport routes of Potosi silver from what is today Bolivia to the then metropolis Spain. Argentina, as most of South America, became an independent country in the early XIXth century. Uncharacteristically, Argentinean Independence was a rather complicated process. It started with the May Revolution of 1810, continued through the July 9th declaration in 1816 (when the United Spanish Provinces of South America declare independence from Spain, unilaterally) and concludes in 1824 with the defeat of the Spanish Empire in the battle of Ayacucho. Characteristically though, the following fifty or so years were marked by severe political instability. There was a long sequence of civil wars, mostly opposing the interests of Buenos Aires (the capital) to those of the provinces (Lynch 1985). Economically, this is a period of modest growth rates which ended with national unification. The Industrial Revolution in Europe fueled demand for primary products and provided new means to satisfy it through important technological innovations: around 1875 the transportation of meat from the other side of the world was made possible and it was made cheap.

There is little disagreement among economists that the period from 1875 to the eve of World War I is the Golden Age, or the Belle Époque, of Argentinean economic history (Taylor, 1992; Sanz-Villarroya, 2007). Just to illustrate this, note that for the year 1913, della Paolera and Taylor (2003) estimate income per capita in Argentina to be (in 1992 US Dollars) around USD 3,797. They provide evidence that this figure is higher than the corresponding figures for France and Germany (USD 3,452 and USD 3,134, respectively) and is substantially higher than those for Spain or Italy. Massive inflows of foreign capital (physical as well as human) supported the rapid expansion of the exports of primary products (grain, meat, wool and leather) which couple with favorable international conditions, ultimately fuelled very rapid rates of economic growth (Rock, 1985, Cortes Conde, 2009). There is also little disagreement that the Argentina's uniqueness is because no other country climbed down so dramatically from the selected group of advanced, rich or developed countries.

The major disagreement among economic historians to this day is not whether but actually when (and, of course, why) this unchecked decline started. Some argue that it started with the 1930 crisis (e.g., Diaz-Alejandro 1985), others argue for an earlier turning point (for instance, Taylor suggests 1913), while Sanz-Villarroya (2005) estimates that the first important structural break for Argentina happens in 1899.[33]

[33] Below we present and discuss our Bai-Perron estimates of the date of structural breaks in Argentinean growth. We find (and adjust our estimates accordingly below) evidence for two structural breaks: 1922 and 1964 (for a full treatment of this issue, see Campos et al. 2008).

Irrespective of exactly when the decline started, its existence was not undisputed until the immediate post II World War. In 1947 Argentina was still ranked the 10th country in the world in terms of per capita income (Alston and Gallo, 2007, p. 6). della Paolera and Taylor (2003) note that "by 1900 Argentina's income per capita had risen from about 67 per cent of developed country-levels in 1870, to 90 percent in 1900, and 100 per cent in 1913 Whatever its exact status in 1913, for all practical purposes Argentina was an advance country" (2003, p. 5). They also calculate that since then the ratio of Argentina's income to OECD income fell to 84 percent in 1950, then to 65 percent in 1973, and then to 43 percent in 1987. This ratio rebounds in the 1990s but again reverts with the 2001 crisis.[34] Last, but not least, it should not go unnoticed that in a recent book on the Great Depressions of the XXth Century (Kehoe and Prescott, 2007), Argentina is the only country that has two chapters (out of 16) entirely and solely dedicated to its economy.

It is not surprising, therefore, that there is a vast literature on the Argentine puzzle, providing alternative explanations for its long-run relative economic decline. One argument is that increased direct competition in international markets during and after WWI (especially from the other areas of new settlement, i.e. Australia and Canada) has an important role to play, as does the sharp decline in immigration and foreign capital inflows. One other argument is that its relative decline is well explained by the fact that the agricultural frontier was reached much earlier in Argentina than in Australia and Canada. Australia's restrictive immigration policy contrasts with Argentina's liberal one, which has been blamed by Diaz-Alejandro (1985) among others, for the difficulties in sustaining and raising productivity levels. Solberg (1987) argues for another reason, this time in terms of Argentina's adoption of a land distribution policy that favored large farm holdings and sustained high levels of wealth inequality. In light of the very accommodating migration policy, the large inflows of workers end up concentrating in Buenos Aires and gave rise to a well-organized and increasingly powerful worker's union movement. It

[34] Growth was negative from 1999 onwards culminating with around -10% in year 2002. The 2001 crisis entailed a default on large part of the external debt, devaluation, inflation, and the freezing of bank accounts (the corralito.) Riots, looting and anti-government demonstrations followed. See Kehoe (2003) for a discussion.

suffices to say that this movement was intimately tangled with the Peron governments, after WWII.

Finance has also received a great deal of attention in terms of its potential role in explaining the Argentinean decline (della Paolera and Taylor, 1998). For example, Prados de la Escosura and Sanz-Villarroya (2006) argue that contract intensive money is actually the key factor in explaining the Argentinean puzzle. Taylor (2003) associates the Argentine decline to extremely low savings rates (the high population dependency rate linked to the immigration policy). This argument combines with Solberg's view and highlights the issue of (restricted) access to finance as a way of perpetuating high inequality levels. Moreover, the role of the financial sector does not need to be limited to domestic or national aspects. Many believe that there may have been excessive dependence on foreign capital in the Belle Époque (British foreign capital to be precise) and the associated radical changes around WWI as an important cause of the Argentinean decline (Taylor, 1992).

Such radical shifts in market conditions extended from the financial to the goods markets, the emphasis here being on international trade. Until 1914, Argentina was an aggressive exporter exhibiting extremely high levels of openness to international trade (measured as the ratio of exports plus imports to GDP.) The data we use in this paper (more details below) shows that this ratio exceeds 50% in the years immediately before WWI, with a clearly declining trend in the inter-wars years (the ratio goes down from about 45% to 20% in these twenty years), and it never exceeds 25% from 1945 to almost 2000. If one believes that exports alone are a major driving force of economic growth, then these numbers surely provide fuel to placing openness as a major reason for the Argentine decline (Diaz-Alejandro, 1970). One important caveat that should be mentioned in this context is that it is unclear (and still much debated) what were the reasons for such a reversal. In particular, the debate is whether this was mainly the disruption and closing up of international markets first with WWI and then with the Great Depression, or was it mainly the adoption of excessively protectionist policies by successive Argentinean governments. Note that these policies inspired and were later reinforced by the import substitution model advocated by the leading Latin American economist of the time, Raul Prebish (from Argentina.)

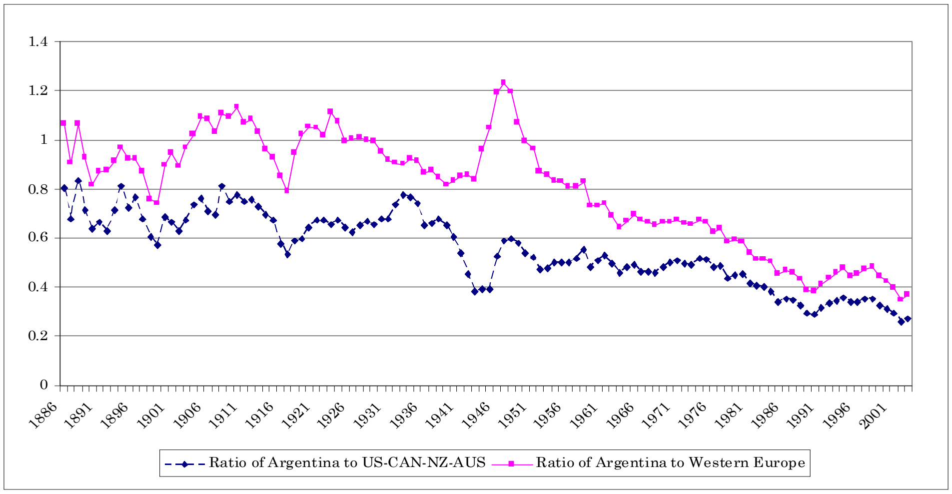
In addition to trade policies, many scholars believe that standard macroeconomic policies, in general, and their inconsistency and the resulting macroeconomic instability, in particular, are also to blame. For instance, della Paollera et al. (2003) show how public deficits throughout Argentinean history (and inflation, mostly since the 1970s), also seem to play an important role in explaining the decline.

Although there is a large literature associating the long-run relative decline of the Argentinean economy with political and institutional factors,[35] we are unaware of studies that try to quantitatively evaluate this association. For instance, Acemoglu and Robinson (2006) observe that: “The political history of Argentina reveals an extraordinary pattern where democracy was created in 1912, undermined in 1930, re-created in 1946, undermined in 1955, fully re-created in 1973, undermined in 1976, and finally reestablished in 1983” (2006, p. 7). In a recent paper, Alston and Gallo (2007) identify the onset of widespread electoral fraud in the 1930s as a turning point for the erosion of the rule of law and one main reason for the Argentinean decline.

In what follows, we take these considerations on board in trying to provide a comprehensive quantitative account of the relative importance of the main reasons often identified with the Argentinean debacle, namely political instability, domestic financial development, trade openness, macroeconomic volatility (inflation and public deficits) and integration in the international financial system.

35 See also della Paolera and Taylor (2003) and references therein.

Fig. 2.8. Ratio of Argentina's GDP per Capita to Developed Countries' GDP per Capita, 1885-2003



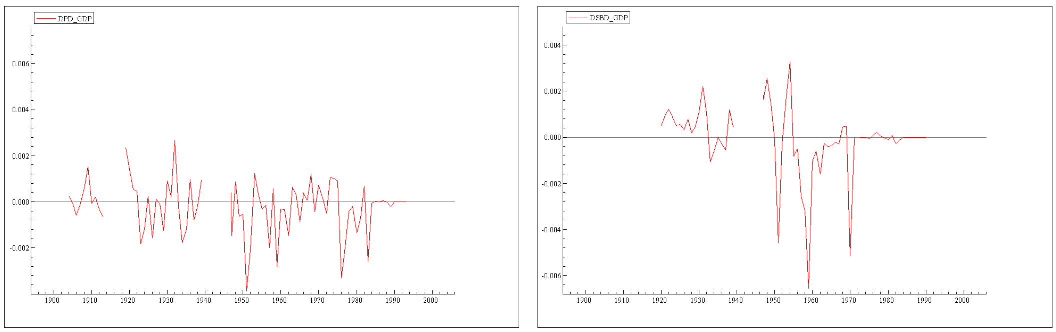
2.3 Data

The data set we put together for this paper reflects the main factors identified by economic historians discussed above. The factors often associated with the relative economic decline of Argentina are the following: financial development, political instability (or institutions), macroeconomic volatility, inflation, trade openness, public deficit, and international financial integration.

Our basic data source is the Cross National Time Series Data set (Banks 2005) which contains historical series on income per capita and various dimensions of instability.[36] This is a commercial database that has been extensively used in the scholarship on growth and political instability (Durlauf et al., 2005.) Data are available yearly for Argentina from 1896 until 2000, for various instability series, excluding the two World War years (that is, 1914 to 1918 and 1939 to 1945).

[36] We have obtained GDP growth and level figures from various other sources (as well as industrial output series) and initial results (not reported) show that these different measures do not affect our results below.

Fig. 2.9. Measures of Financial Development



Private deposits over GDP   Savings banks deposits over GDP

Our two main measures of financial development try to capture the efficiency of the financial sector, not its relative size. The source for both is Mitchell (2003). The first is the bank deposits by the private sector over GDP (private deposits), which we believe is a good proxy for the share of credit to the private sector over GDP. Our second measure from Mitchell (2003) is the total deposits in savings banks over GDP. Given its more restrictive nature and the fact that the exact definition of savings bank deposits contains an unobservable legal element, we use this variable mostly for robustness check thereby attaching greater weight to private deposits.[37]

We also explore the hypothesis that different types of political instability have different effects on economic growth.[38] This is done by further developing the distinction between formal and informal political instability introduced in Campos and Karanasos (2008). The distinction is based on whether or not different forms of instability originate from

---

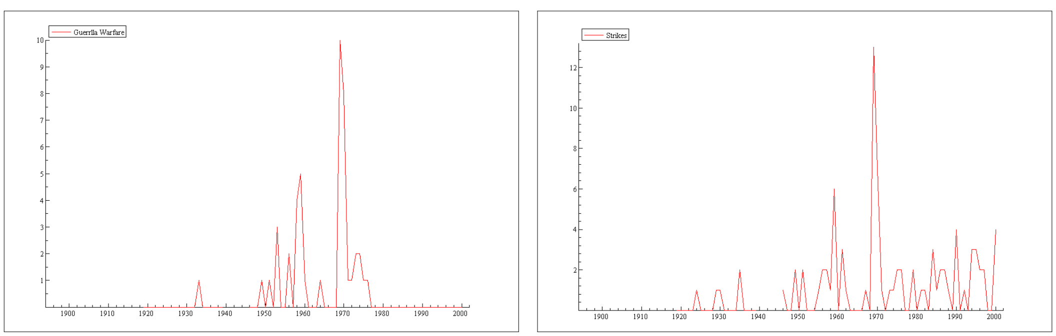
37 For robustness we also use two measures of financial development that reflect depth. The first indicator we use is the ratio of M3 to GDP, from Alston and Gallo (2007). The main reason for considering this measure is that it has been used extensively in the finance-growth literature (see Campos et al. 2011). We also use a narrower version of this variable (M1 over GDP) to further check for the robustness of our results (source of data is Bordo et al., 2001). For the Figures and the results see the Appendix in Campos et al. (2011a).

38 Another puzzle we are interested in regards the duration of the political instability effects: while the conventional wisdom is that these are severe in the long-run, placeCityCampos and Nugent (2002) and Murdoch and Sandler (2004) argue that they are significantly stronger in the shorter- than in the long-run. In placeCity-Campos and Nugent (2002), the long-run effect vanishes when the African countries are excluded from the estimation and when institutions are taken into account.

within the political system: guerrilla warfare are thus informal political instability, while constitutional reforms are classified as formal instability. In addition to the obvious policy implications this taxonomy generates (in a literature in which policy implications are scarce), this distinction allows us to investigate questions that naturally have not been investigated so far, such as whether or not the effects of some forms of informal instability are more severe in the short- than in the long-run, and whether or not the main effect of formal instability occurs through growth volatility. One of our hypotheses is that the answer to these questions is the same ("yes") and below we provide further justification as well as full econometric support.

Our informal political instability variables[39] are strikes (a general strike of 1,000 or more workers involving more multiple employers and aimed at government policies) and guerrilla warfare (armed activity, sabotage, or bombings by independent bands of citizens and aimed at regime overthrow). These series are available since 1919 (Figure 2.10).

Fig. 2.10. Measures of Informal Political Instability



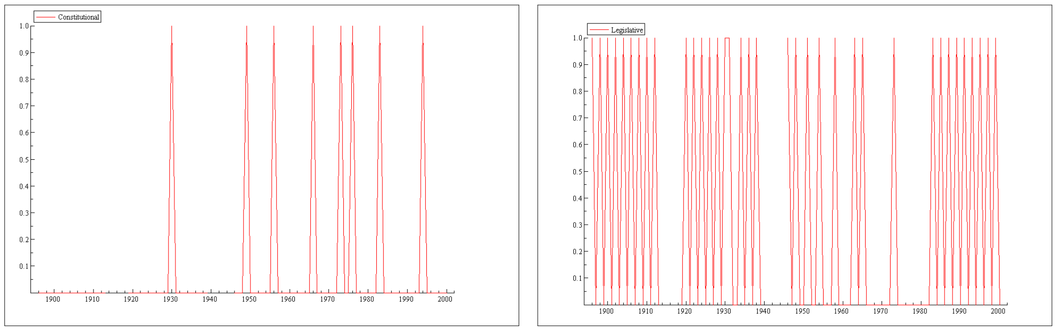
Guerilla Warfare　　　　Strikes

Our formal political instability variables (Figure 2.11) are as follows: the number of constitutional changes and legislative elections.[40]

[39] Our political instability variables enter one by one in the econometric framework we use, so our results are not affected by the taxonomy and as such it is used simply to facilitate the interpretation.

[40] For robustness we use two more measures of informal political instability: annual number of anti-government demonstrations (peaceful public gatherings of at least 100 people) and assassinations (defined

Fig. 2.11. Measures of Formal Political Instability



Constitutional changes Legislative elections

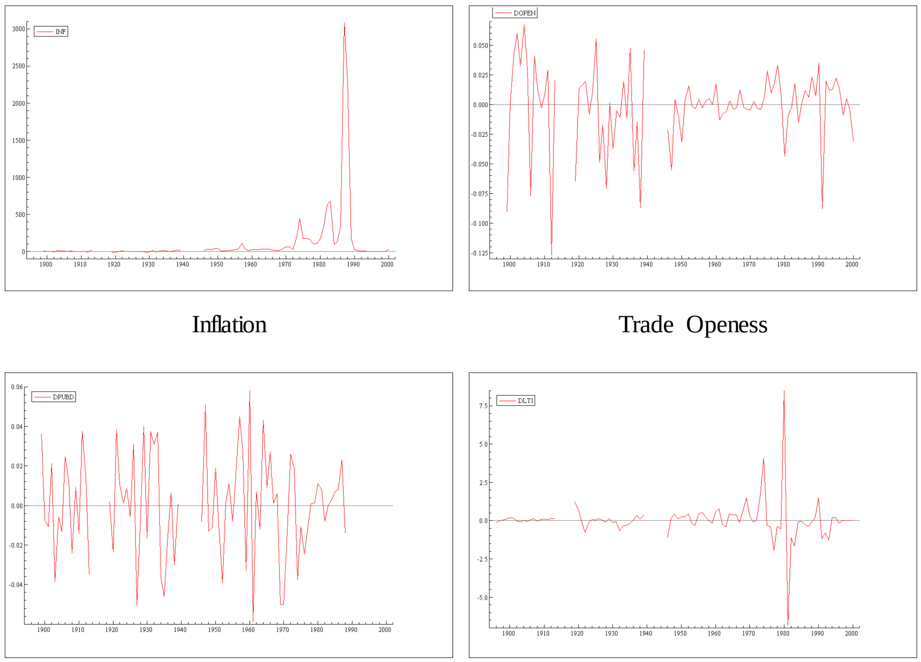
Our measures of inflation, trade openness and public deficit are from Alston and Gallo (2007). Inflation is measured as yearly changes in the consumer price index (CPI). Public deficit is proxied as the ratio of the federal deficit to GDP, but it does exclude state-owned enterprises.[41] Trade openness is measured in standard fashion as the ratio of imports plus exports to GDP. Alston and Gallo (2007) have carried out various necessary adjustments to underlying data from Venganzones and Winograd (1997), from the Ministry of Economy of Argentina and from the IMF's International Financial Statistics.

Finally, international financial sector developments have also been repeatedly blamed for Argentina's poor economic performance. There are two aspects of this issue that are often said to play a role: the first being the credit crunch associated with the onset of WWI and with the Great Crisis of 1929, and the second being the change in global financial leadership which went from London to New York during this period. We must say that we proceed as if the second aspect is less important, but also that we are absolutely sure it is much more difficult to measure than the first. Thus, in standard fashion in this type

as politically motivated murders or attempted murders of a high government official or politician). We also use two more measures of formal political instability: the number of cabinet changes and the size of the cabinet (see Figure 2.14 in Appendix).

41 Because the original inflation series contain a number of obvious outliers between the years 1987 and 1991 (reaching almost 5,000% in 1989), we lower the relative weight of these observations for estimation.

Fig. 2.12. Other Variables



Inflation — Trade Openess

Public Deficit — UK Interest Rate

of study, we use the level of interest rates in the United Kingdom as our proxy for the overall conditions in international financial markets (the source of these data is Bordo et al. 2001). Because the transition to the U.S. financial leadership is often said to be even less beneficial to Argentina (mainly because American investors often refrained to take managerial control of Argentine firms), our estimates for this effect should be conservative and if at all biased will show a smaller than actual effect of the international financial market in the Argentinean decline.

2.4 Econometric Framework

The PARCH model was introduced by Ding, Granger and Engle (1993) and quickly gained currency in the finance literature.[42] Let growth (y_t) follow a white noise process

[42] See, for example, Karanasos and Kim (2006). Karanasos and Schurer, (2005, 2008) use this process to

augmented by a risk premium defined in terms of volatility:

$$y_t = c + kh_t + \lambda x_{it} + \epsilon_t, \tag{2.10}$$

with

$$\epsilon_t = e_t h_t^{\frac{1}{2}},$$

where x_{it} is either the political instability or the financial development variable or one of the other explanatory variables.[43]

In addition, $\{e_t\}$ are independently and identically distributed (i.i.d) random variables with $E(e_t) = E(e_t^2 - 1) = 0$, while h_t is positive with probability one and is a measurable function of the sigma-algebra $\sum_{t-1}$, which is generated by $\{y_{t-1}, y_{t-2}, \ldots\}$.

In other words, h_t denotes the conditional variance of growth. In particular, h_t is specified as an asymmetric PARCH(1,1) process with lagged growth included in the variance equation:

$$h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} f\left(e_{t-1}\right) + \beta h_{t-1}^{\frac{\delta}{2}} + \phi x_{it} + \gamma y_{t-l}, \tag{2.11}$$

with

$$f\left(e_{t-1}\right) = \left[\left|e_{t-1}\right| - \varsigma e_{t-1}\right]^{\delta},$$

where δ (with $\delta > 0$) is the heteroscedasticity parameter, α and β are the ARCH and GARCH coefficients respectively, ς with $|\varsigma| < 1$ is the leverage term and γ is the level term for the lth lag of growth.[44] In order to distinguish the general PARCH model from

---

model output growth and inflation respectively.

[43] Because the original financial development, openness, public deficit and country-regionplace UK interest rate variables, are I(1), they enter our models in first differences.

[44] The model imposes a Box-Cox power transformation of the conditional standard deviation process and the asymmetric absolute residuals.

a version in which δ is fixed (but not necessarily equal to two) we refer to the latter as (P)ARCH.

The PARCH model increases the flexibility of the conditional variance specification by allowing the data to determine the power of growth for which the predictable structure in the volatility pattern is the strongest. This feature in the volatility process has important implications for the relationship between political instability, finance, inflation, and growth and its volatility. There is no strong reason for assuming that the conditional variance is a linear function of lagged squared errors. The common use of a squared term in this role is most likely to be a reflection of the normality assumption traditionally invoked. However, if we accept that growth data are very likely to have a non-normal error distribution, then the superiority of a squared term is unwarranted and other power transformations may be more appropriate.

The PARCH model may also be viewed as a standard GARCH model for observations that have been changed by a sign-preserving power transformation implied by a (modified) PARCH parameterization. He and Teräsvirta (1999) emphasize that if the standard Bollerslev type of model is augmented by the heteroscedasticity parameter (the power term), the estimates of the ARCH and GARCH coefficients almost certainly change.[45]

We present our main reasons in three interdependent blocs: the direct, indirect and dynamic (short and long-run) effects. We proceed with the estimation of the PARCH(1,1) model in equations (2.10) and (2.11) in order to take into account the serial correlation observed in the levels and power transformations of our time series data. The tables below report the estimated parameters of interest for the period 1896-2000. These were obtained by quasi-maximum likelihood estimation (QMLE) as implemented in EVIEWS. The best fitting specification is chosen according to the Likelihood Ratio (LR) results and the minimum value of the Information Criteria (IC) (not reported). Once heteroscedas-

[45] Karanasos and Schurer (2008) find that the relationship between the variable and its conditional variance is sensitive to changes in the values of the heteroscedasticity parameter. Put differently, the estimated values of the in-mean and the level effects are fragile to changes in the power term.

ticity has been accounted for, our specifications appear to capture the serial correlation in the power transformed growth series.[46]

Our set of variables tries to reflect the different explanations for the Argentinean puzzle previously put forward by economic historians. This set comprises domestic and international financial developments, informal and formal political instability, inflation and public deficit, and the degree of openness to international trade. In order to study the direct effects of our set of explanatory variables, we specify model 1 with $\phi = \gamma = 0$ in equation (2.11), while model 2 with $\lambda = 0$ in equation (2.10) allows us to investigate their indirect impacts on growth. [47]

2.5 Empirical Results

Our results are presented following specific types of effects. That is, we discuss direct (on mean economic growth), indirect (via volatility), dynamic (short and long-run) and structural break effects. Moreover, in trying to satisfy both the time-series and economic growth literature traditions (the former mostly univariate and the latter multivariate), for each effect we report estimates for one variable at a time before discussing the full multivariate results.

2.5.1 Direct Growth Effects

Table 2.13 reports the results from our estimation of the (P)ARCH(1,1) model for each one of the elements in our set of explanatory variables. The parameter we are most interested in is λ (in the third column.) The results reveal that the direct effect of financial

[46] For all cases, we find that the leverage term is insignificant, so we re-estimate **our models** excluding this parameter.

[47] As indicated in the previous section we jointly estimate the conditional mean and variance of growth in order to take into account PARCH effects. At the same time, with a limited number of time-series observations the non-linear structure should not be overextended as this imposes excessive requirements on the data. Therefore, we estimate the direct (model 1) and indirect (model 2) effects separately.

Table 2.13. Direct Effects on Economic Growth: (P)ARCH estimates, Argentina 1896-2003

$x_{it} \downarrow$	k	λ	α	β	δ
Inflation	1.05 (2.42)	-4×10^{-5} (0.71)	0.48 (4.34)	0.69 (7.66)	0.80 –
Trade Openness	0.90 (1.95)	-0.060 (1.73)	0.77 (3.63)	0.47 (2.28)	0.80 –
Public Deficit	0.72 (1.94)	-0.070 (3.13)	0.95 (3.05)	0.43 (2.00)	0.90 –
UK Interest Rate	0.94 (3.86)	-0.001 (0.37)	0.82 (3.64)	0.44 (1.61)	0.90 –
Guerilla Warfare	1.00 (3.69)	-0.001 (4.35)	0.77 (5.43)	0.47 (3.13)	0.90 –
Constitutional Changes	1.80 (1.99)	-0.003 (1.35)	0.56 (3.01)	0.48 (1.25)	0.80 –
Private Deposits/GDP	0.76 (2.66)	0.980 (9.21)	0.70 (4.99)	0.57 (4.94)	0.80 –

Note: This table reports parameter estimates for the following model:

$$y_t = c + kh_t + \lambda x_{it} + \varepsilon_t, h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \mid e_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}}$$

The numbers in parentheses are absolute t statistics.

development (private deposits) on per capita economic growth rates is positive and statistically significant, those of informal political instability (guerrilla warfare), trade openness, and public deficit are negative,[48] whereas the effects of formal political instability, international financial development (interest rate in the United Kingdom) and inflation are not statistically significant, at conventional levels. [49]

As for the in-mean parameter (k), notice that in all cases the estimates are statistically significant and positive which is in line with the theoretical argument of Black (1987). Also the power term coefficients δ are rather stable, with the Akaike IC (AIC) criteria choosing a (P)ARCH specification with power term ranging from 0.8 (e.g., inflation) to 0.9 (e.g., public deficit.) [50]

---

[48] This result for trade openness is clearly unexpected. Notice, however, that we show below that its short-run effect is negative but the long-run impact is positive (see Table 2.17 in below).

[49] We also estimate bivariate regressions (results not reported) to examine the joint effect of informal political instability and financial development on growth. It appears that anti-government demos, assasinations and M3/GDP have no impact on growth. Therefore, in what follows we only use guerrilla warfare, strikes and financial efficiency. The results are not reported for the sake of space. They are available in the Appendix.

[50] Notice that in all our estimations the ARCH and GARCH parameters (α and β) are highly significant in the majority of the cases. Also the estimated power term coefficients are stable ranging from 0.80 to 1.10.

Table 2.14. Direct Effect of Guerilla Warfare/Strikes, Private Deposits/Savings Bank Deposits, UK Interest Rate, and Trade Openness on Economic Growth. (P)ARCH estimates

	k	λ_1	λ_2	λ_3	λ_4	α	β	δ
Guerilla Warfare								
Private Deposits/GDP	0.81 (2.06)	−0.001 (3.32)	0.328 (1.98)	−0.001 (1.40)	−0.011 (1.18)	0.81 (4.42)	0.51 (4.31)	1.00 –
Savings Bank Deposits/GDP	0.75 (1.54)	−0.001 (2.63)	0.147 (2.16)	−0.004 (7.18)	−0.014 (2.50)	0.96 (5.07)	0.52 (7.46)	1.10 –
Strikes								
Private Deposits/GDP	0.74 (2.15)	−0.002 (2.14)	0.263 (1.69)	−0.001 (1.50)	−0.046 (6.89)	0.89 (4.60)	0.43 (2.86)	0.80 –
Savings Bank Deposits/GDP	0.73 (1.62)	−0.001 (2.01)	0.308 (1.79)	−0.004 (5.17)	−0.127 (1.46)	1.00 (3.87)	0.51 (5.05)	1.10 –

This table reports parameter estimates for the following model:

$$y_t = c + kh_t + \lambda_1 x_{it}^{(pi)} + \lambda_2 x_{it}^{(fd)} + \lambda_3 x_{uk,t} + \lambda_4 x_{to,t} + \varepsilon_t,\ h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \mid e_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}},$$

where $x_{it}^{(pi)}$ is either guerilla warfare or strikes, $x_{it}^{(fd)}$ is either private deposits/GDP or savings bank deposits/GDP, $x_{uk,t}$ is UK interest rate, and $x_{to,t}$ is trade openess.

The numbers in parentheses are absolute t statistics.

How robust are these baseline individual results? As discussed above, previous research on Argentina has given considerable weight to the roles of informal political instability and financial development. It is not as surprising therefore that these are the two dominant influences since they are the ones we estimate with greatest precision (more significant). One robustness test would be to investigate whether or not such powerful and precise effects obtain in the presence of the other explanatory variables. In other words, we want to be sure that they remain if we add to the baseline specification any of our four additional variables. [51]

Table 2.14 presents the results when we add all four regressors. That is, informal political instability, domestic and international financial development, and trade openness. Both guerilla warfare and strikes still show the expected negative and statistically significant direct impact (see the λ_1 column). As for the effect of financial efficiency, it is still

[51] Our bivariate/trivariate analysis (results not reported) shows that the direct effect of financial efficiency is not affected by the addition of any of the four explanatory variables to the model, with both indicators still showing a positive and significant effect and the same conclusion remains for the case of guerilla warfare and government strikes (see Tables 2.24-2.29 in Appendix). Interestingly, the direct negative effect of public deficit disappears when we control for either informal political instability or private deposits. The results also reinforce that inflation has no direct impact on growth (see Tables 2.26-2.27 in Appendix).

positive and statistically significant (see the λ_2 column). It is worth noting that the influences of the UK interest rate and the trade openness on growth are qualitatively altered by the presence of informal political instability and financial efficiency. In particular, the significance of their influence is altered by changes in the choice of the informal political instability and financial efficiency variables. More specifically, the negative impact of UK interest rate on growth remains when we include in the model the impact of savings bank deposits on growth but it disappears when we include private deposits. Similarly, trade openness affects growth negatively only in two out of the four cases. (see the λ_3 and λ_4 columns in Table 2.14).[52]

In summary, we find that the main explanatory factors, solely in terms of their direct effects on economic growth in Argentina, turn out to be domestic (mainly financial efficiency) and international (UK interest rate) financial development, and trade openness. Although informal political instability also seems to play an important role, we find that only one aspect seems to survive our battery of sensitivity tests, that being guerrilla warfare and to a lesser extend strikes. We now turn to the investigation of the indirect effects.

2.5.2 Indirect Effects (Via Growth Volatility)

One of the main advantages of the (P)ARCH framework is that it allow us to study not only the direct growth effects from the full set of explanatory variables described above, but also their indirect effects on economic growth through the predicted component of growth volatility (conditional on its past values). As we can see from Tables 2.13 and 2.14 above and from Tables 2.15 and 2.16 in this section, the effect of conditional or predicted volatility on growth is in all cases positive ($k > 0$) and statistically significant at conventional levels. By construction, an increase in the unexpected volatility should decrease economic growth rates?. In the current section, we present our results for such

[52] Interestingly, this is not the case when we use a trivariate analysis. That is, when we include as a regressor either UK interest rate or trade openess (but not both) then we find the expected negative and significant influence in all four cases (see Table 2.28 in Appendix). Finally, as with the bivariate analysis in the trivariate one, public deficit and inflation have no effect on growth (see Table 2.29 in Appendix).

indirect effects in two parts and follow the same format as before: we first discuss the indirect effects of each one of our explanatory variables and then we present results for our complete set (that is, including all the main explanatory variables).

Table 2.15 reports the estimation results for each one of the elements in our data set for what we call the indirect effect, which is the effect on growth via the volatility channel.[53] The parameter we are most interested in is ϕ (in the fifth column.) Our results show that the indirect effects of trade openness and formal political instability (in this case, constitutional changes) on the conditional volatility of per capita economic growth rates are negative and statistically significant whereas those of inflation, public deficit and UK interest rate are positive and significant. Interestingly, the volatility of growth is independent of changes in financial development and formal political instability.

Larger public deficits, escalating inflation rates and (to a lesser extent) higher UK interest rates are associated with a larger fraction of growth volatility that is anticipated by the relevant economic agents. And the larger the share of the total growth volatility that is anticipated, the higher the per capita growth rates we observe. Therefore, inflation, UK interest rate and public deficits register a negative direct effect on growth (the first two an insignificant one) but a positive and substantial impact on the expected or conditional share of growth volatility (see Tables 2.13 and 2.15).

On the other hand, we find that exogenous increases in trade openness has a negative and significant impact on the conditional growth volatility (recall that the direct effect is also negative). This result reflects one of the costs many economists associate with trade liberalization efforts: in the short-run, changes in the share of trade in GDP decrease the conditional or expected share of growth volatility (or, equivalently, increase the amount of growth volatility that economic agents are not able to anticipate.) Therefore such a decrease in conditional volatility driven by trade openness translate into lower rates of economic growth (because $k > 0$).

[53] In the expressions for the conditional variances reported in the tables, various lags of growth (from 1 to 12) were considered with the best model (l = 6) chosen on the basis of the minimum value of the AIC.

Last, and also of interest, is that we could not detect any significant indirect effects from domestic financial development (proxied by private deposits) or informal political instability (proxied by the occurrence of guerilla warfare). There is no evidence that such factors affect growth in Argentina indirectly, through the conditional volatility of growth. Recall, however, that we do find that the direct effects of both domestic financial development and guerrilla warfare are substantial (see Table 2.14).

It is also worth noting that since the estimates for the in-mean parameter (k) and the level coefficient (γ) in Table 2.15 are statistically significant and positive they offer strong strong evidence for a positive bidirectional feedback relationship between growth and its volatility,[54] which seems robust to the presence of various finance and instability variables.

We now proceed by investigating the robustness of these results. Specifically, and for comparability purposes, we ask how the results from the various aspects of financial development and political instability change if we add to the baseline model the complete set of explanatory variables (as opposed to assess their effects one by one).[55]

Table 2.16 shows that after adding this full set of controls, the indirect negative effect of formal political instability remains statistically significant throughout. Focusing our attention first on the ϕ_1 and k parameters, note that both forms of formal instability (in this case, the occurrence of changes in the constitution or of legislative elections in a given year) are found to affect conditional volatility negatively ϕ_1 $(\phi_1 < 0)$. Since $k > 0$, con-

---

[54] The existing empirical literature focuses mainly on the effect of volatility on growth, see Fountas et al. (2006), Fountas and Karanasos (2007), Conrad et al. (2010) and Conrad and Karanasos (2011).

[55] For the sake of space, we do not report the results for the intermediate steps (namely, between the results for one by one variables and for all variables together). These results for the indirect effects for each pair of variables and each three are thus available upon request. When we include in the variance of growth the UK interest rate and/or trade openess, and one of the four alternative measures of formal political instability the effects of cabinet changes and cabinet size disappear (see tables 2.30-2.31 and 2.34 in Appendix). Therefore, in what follows we only use constitutional changes and legislative elections. Moreover when we control for formal political instability inflation has no impact on growth volatility (see Table 2.33 in Appendix). In addition, our bivariate/trivariate analysis shows that the effects of the UK interest rate, trade openess and public deficit are not affected by the addition of any of the four measures of formal political instability (see tables 2.30-2.32 and 2.34 in Appendix).

Table 2.15. Indirect effect on Economic Growth. (P)ARCH estimates

$x_{it} \downarrow$	k	α	β	ϕ	γ	δ
Inflation	1.66 (3.03)	0.60 (5.36	0.54 (7.52)	5×10^{-5} (2.98)	0.24 (2.44)	0.80 –
Trade Openness	1.65 (2.72)	0.75 (6.41)	0.28 (0.16)	−0.200 (3.48)	0.19 (2.69)	0.80 –
Public Deficit	1.02 (2.73)	0.69 (5.45)	0.45 (5.15)	0.120 (2.44)	0.29 (4.76)	0.80 –
UK Interest Rate	1.55 (2.26)	0.50 (2.98)	0.55 (5.35)	0.004 (1.57)	0.15 (4.67)	1.00 –
Guerilla Warfare	1.12 (2.46)	0.73 (4.80)	0.46 (4.00)	0.001 (0.82)	0.10 (2.00)	0.90 –
Constitutional Changes	1.18 (1.94)	0.69 (4.40)	0.45 (4.15)	−0.008 (3.40)	0.18 (3.75)	1.00 –
Private Deposits/GDP	2.05 (2.23)	0.41 (3.04)	0.62 (6.75)	0.580 (0.53)	0.40 (5.69)	0.80 –

Note: This table reports parameter estimates for the following model:

$$y_t = c + kh_t + \varepsilon_t,\ h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \mid e_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}} + \phi x_{it} + \gamma y_{t-6}$$

The numbers in parentheses are absolute t statistics.

stitutional changes affect growth negatively as well. Economic agents have severe difficulties in anticipating the consequences of changes in the rules of the game (constitutions) and in the composition of the legislature following democratic elections. Such changes increase the share of unanticipated uncertainty and this accordingly reduces growth. Of course, these results reinforce the notion that the type of political instability matters vis-à-vis economic growth: while informal may have a direct effect, the impact of formal instability operates indirectly, via growth volatility.

There a number of additional important results from Table 2.16. In terms of the effects of financial globalization (or, more specifically, of the international dimensions of financial development), we find that they tend to be positive and significant $(\phi_2 > 0)$ on anticipated growth volatility when proxied by the interest rate in the United Kingdom. This is intuitive as reductions in the UK interest rate translate into the reduction of the price of money internationally with the latter pricing accounting for risk. Moreover, the impact of public deficit on volatility also remains positive $(\phi_4 > 0)$ and statistically significant. On the other hand, we find evidence that increases in trade openness are associated with decreases in conditional volatility $(\phi_3 < 0)$ of per capita growth in Argentina.

Table 2.16. Indirect Effect of Constitutional Changes/Legislative Elections, UK Interest Rate, Trade Openness, and Public Deficit on Economic Growth. (P)ARCH estimates

	k	α	β	ϕ_1	ϕ_2	ϕ_3	ϕ_4	γ	δ
Constitutional Changes	1.52 (3.09)	0.46 (4.04)	0.69 (7.03)	-0.02 (6.19)	0.01 (2.99)	-0.17 (5.91)	0.08 (2.96)	0.04 (0.74) $_{l=7}$	1.00 –
Legislative Elections	5.24 (1.39)	0.13 (0.76)	0.96 (7.72)	-0.01 (1.74)	0.01 (2.43)	-0.11 (2.35)	0.12 (3.69)	-0.06 (0.49) $_{l=6}$	0.90 –

Note: This table reports parameter estimates for the following model: $y_t = c + kh_t + \varepsilon_t$,
$h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \mid e_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}} + \phi_1 x_{it}^{(pi)} + \phi_2 x_{uk,t} + \phi_3 x_{to,t} + \phi_4 x_{pd,t} + \gamma y_{t-l}$,
where $x_{it}^{(pi)}$ indicates either constitutional changes or legislative elections,
$x_{uk,t}$ is UK interest rate, $x_{to,t}$ is trade openness, and $x_{pd,t}$ is public deficit.
The numbers in parentheses are absolute t statistics.

In summary, we find strong evidence that both informal political instability (constitutional changes) and trade openness have a negative indirect (via volatility) impact on growth whereas UK interest rate and public deficit affect it positively. No other variables in our set of explanatory variables seem to exhibit equally robust estimates of their indirect effects.

2.5.3 Dynamic Aspects

In this section we investigate how short- and long-run considerations help us refine our baseline results. Another potential benefit from this exercise is that the required use of lags may help ameliorate any lingering concerns about endogeneity. In order to estimate short- and long- run relationships we employ the following error correction (P)ARCH form

$$\Delta y_t = \mu + \theta \Delta x_{i,t-l} + \varphi(y_{t-1} - c - \zeta x_{i,t-1}) + \varepsilon_t, \tag{2.12}$$

where θ and ζ capture the short and long-run effects respectively, and φ is the speed of adjustment to the long-run relationship.[56] This is accomplished by embedding a long-

[56] As pointed out by Loayaza and Rancière (2006) the requirements for the validity of this methodology are that: i) there exists a long-run relationship between the variables of interest and, ii) the dynamic specification of the model is sufficiently augmented so that the regressors are strictly exogenous and the resulting residual is serially uncorrelated.

run growth regression into an ARDL model.[57] In other words, the term in parenthesis contains the long-run growth regression, which acts as a forcing equilibrium condition

$$y_t = c + \zeta x_{it} + u_t, \tag{2.13}$$

where ϵ_t is $I(0)$. The lag of the first difference of either the political instability or financial development variable or one of the explanatory variables ($\Delta x_{i,t-l}$) characterizes the short-run effect. The condition for the existence of a long-run relationship (dynamic stability) requires that the coefficient on the error-correction term be negative and not lower than -2 (that is, $-2 < \varphi < 0$). We also take into account the PARCH effects by specifying the error term u_t as follows

$$\varepsilon_t = e_t h_t^{\frac{1}{2}}, \tag{2.14}$$

where

$$h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \left| e_{t-1} \right|^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}}. \tag{2.15}$$

Table 2.17 presents the results on the estimation of short and long-run parameters linking the four explanatory variables with growth. In all cases, the estimated coefficient on the error correction term (φ) lies within the dynamically stable range $(-2, 0)$. From investigating whether dynamic considerations affect our conclusions, we find important differences in terms of short and long-run behavior of our explanatory variables, more specifically, while the effects (negative) of informal political instability, public deficit and UK interest rate are similar in the long- and short-run, that of the financial development and trade openness are negative in the short- and positive in the long-run (see the θ and ζ columns). Interestingly, the effect of inflation is not statistical significant.

For the sake of space, we do not report the results for the intermediate steps (namely, between the results for one by one variables and for all variables together). These results

[57] For details on the "ARDL approach," see Pesaran (1997) and Pesaran and Shin (1999).

Table 2.17. The short- and long-run effects on Growth

$x_{it} \downarrow$	θ	φ	ζ	α	β	δ
Inflation	-5×10^{-5} (0.75) $l=0$	-0.88 (9.98)	0.0001 (1.53)	0.91 (5.22)	0.42 (4.81)	0.90 –
Trade Openness	-0.1000 (1.45) $l=3$	-0.58 (4.83)	0.2500 (2.01)	0.65 (5.16)	0.61 (7.64)	0.90 –
Public Deficit	-0.1100 (9.08) $l=0$	-0.60 (25.94)	-0.1700 (5.88)	1.42 (4.08)	0.22 (2.16)	0.80 –
UK Interest Rate	-0.0150 (4.77) $l=5$	-0.39 (7.47)	-0.0160 (4.21)	1.27 (3.81)	0.14 (0.95)	0.90 –
Guerilla Warfare	-0.0014 (3.38) $l=3$	-0.60 (7.20)	-0.0007 (2.59)	1.10 (4.19)	0.36 (3.59)	0.90 –
Private Deposits/GDP	-1.3500 (1.81) $l=5$	-0.44 (4.64)	0.9399 (23.72)	0.37 (2.63)	0.80 (6.69)	0.90

Note: This table reports parameter estimates for the following model:
$\Delta y_t = \mu + \theta \Delta x_{i,t-l} + \varphi(y_{t-1} - c - \zeta x_{i,t-1}) + \varepsilon_t$,
$h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \left| e_{t-1} \right|^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}}$. θ (l is the order of the lag)
and ζ capture the short- and long-run effects respectively.
φ indicates the speed of adjustment to the long-run relationship.
The numbers in parentheses are absolute t statistics.

for each pair of variables and each three are thus available upon request.[58] Table 2.18 presents the results when we add all four regressors. That is, informal political instability, domestic and international financial development, and trade openness.

The estimates of φ lie within the range -0.68 to -0.36. Regarding the short and long-run estimates, θ_i and ζ_i we focus our analysis first on those obtained from the informal political instability variables. All four estimates of the short-run coefficients (see the θ_1 column) are highly significant and negative and their absolute values are higher than the corresponding values for the long-run coefficients (see the ζ_1 column). This provides supporting evidence for the notion that the duration of the political instability effect does indeed matter and, for guerilla warfare and general strikes, such effects tend to be considerably stronger in the short- than in the long-run as previously noted by Campos and

[58] The results from the bivariate analysis suggest that from the four informal political instability variables only guerilla warfare and strikes affect significantly growth in the long-run (see Table 2.37 in Appendix). That is when we control for the financial development the long-run effects of anti-government demos and assassinations disappear. Similarly, M3/GDP has no long-run effect on growth in three out of the four cases. Finally, when we control for informal political instability and financial efficiency the effect of public deficit on growth disappears.

Nugent (2002), Murdoch and Sandler (2004), Campos and Karanasos (2008) and Campos et al. (2011). As with the univariate analysis both the short- and long-run effects of the UK interest rate are negative (see the θ_3 and ζ_3 columns).

Next we discuss the results regarding the financial development variables. In the long-run, we find that financial development affects growth positively (see the ζ_2 column). This result is very much in line with the large empirical literature reviewed by Levine (2005) and it is interesting we can reproduce it with our rather different methodology. Maybe more interestingly, the short-run coefficients tell a very differently story: we find that the short-run impact of financial development on growth is negative and significant (see the θ_2 column). Thus our results square well with recent findings by Loayaza and Rancière (2006), among others, in that the sign of the relationship between economic growth and financial development depends on whether the movements are temporary or permanent (the effect being negative in the former and positive in the latter.) It is noteworthy that the stronger long-run effects we obtain are for savings bank deposits. Finally, the long-run impact of trade openness disappears.

In summary, in the short-run all four variables have a negative effect on growth. In the long-run informal political instability and the UK interest rate affect growth negatively whereas the impact of financial efficiency turns to positive and that of trade openness disappears.

Table 2.18. The short- and long-run Growth effects of Guerrila Warfare/Strikes, Private Deposits/ Savings Bank Deposits, UK Interest Rate and Trade Openness

$x_{it}^{(fd)} \downarrow$	θ_1	θ_2	θ_3	θ_4	φ	ζ_1	ζ_2	ζ_3	ζ_4	δ
					Guerrila Warfare					
Private Deposits/GDP	-0.0084 (2.24) $l=2$	-0.1114 (1.69) $l=3$	-0.0020 (4.44) $l=0$	-0.0237 (5.22) $l=0$	-0.6814 (5.69)	-0.0012 (3.65)	0.2348 (1.51)	-0.0034 (3.05)	-0.0084 (1.40)	0.80 –
Savings Bank Deposits/GDP	-0.0018 (3.28) $l=3$	-0.3117 (4.27) $l=3$	-0.0053 (25.48) $l=0$	-0.0058 (3.91) $l=1$	-0.4593 (4.31)	-0.0009 (2.84)	0.1708 (2.21)	-0.0039 (5.54)	-0.0108 (1.43)	0.80 –
					Strikes					
Private Deposits/GDP	-0.0023 (7.70) $l=3$	-0.0949 (2.48) $l=1$	-0.0056 (19.57) $l=0$	-0.0112 (8.41) $l=3$	-0.4516 (9.60)	-0.0010 (2.23)	0.3205 (1.57)	-0.00395 (4.95)	-0.0059 (0.67)	0.80 –
Savings Bank Deposits/GDP	-0.0015 (3.54) $l=3$	-0.2052 (2.02) $l=1$	-0.0025 (5.21) $l=0$	-0.0194 (3.79) $l=0$	-0.3582 (7.13)	-0.0006 (1.95)	0.3779 (3.47)	-0.0038 (5.08)	-0.0092 (1.19)	0.80 –

Note: This table reports parameter (mean) estimates for the following model:

$$\Delta y_t = \mu + \theta_1 \Delta x_{i,t-l}^{(pi)} + \theta_2 \Delta x_{i,t-l}^{(fd)} + \theta_3 \Delta x_{uk,t-l} + \theta_4 \Delta x_{to,t-l} + \varphi(y_{t-1} - c - \zeta_1 x_{i,t-1}^{(pi)} - \zeta_2 x_{i,t-1}^{(fd)} - \zeta_3 x_{uk,t-1} - \zeta_4 x_{to,t-1}) + \varepsilon_t,$$

$h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \left|\varepsilon_{t-1}\right|^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}}$. The four θ's (l is the order of the lag) and ζ's capture the short- and long-run effects respectively.

φ indicates the speed of adjustment to the long-run relationship.

$x_{i,t-l}^{(pi)}$ denotes an informal political instability (either guerilla warfare or strikes) variable.

$x_{i,t-l}^{(fd)}$ denotes a financial development (either private deposits/GDP or savings bank deposits/GDP) variable.

$x_{uk,t-l}$ is UK interest rate, $x_{to,t-l}$ is trade openness.

The numbers in parentheses are absolute t statistics.

2.5.4 Structural Breaks

One final important robustness test regards the role of structural breaks. We use the methodology developed by Bai and Perron (2003) to examine whether there are any structural breaks in growth, its volatility, the various political instability series and the first differences of the four financial development variables. Bai and Perron (2003) address the problem of testing for multiple structural changes under very general conditions on the data and the errors. In addition to testing for the existence of breaks, these statistics identify the number and location of multiple breaks.

In the case of the economic growth series (and, interestingly, also for growth volatility) the Bai-Perron methodology supports two structural break points.[59]The first occurs for year 1922 and the second for year 1964. For our political instability variables, we find no structural breaks for the guerilla warfare and constitutional changes series,[60] and we also find no breaks in the four financial development variables.

However, our Bai-Perron results support that general strikes have one structural break, which is dated for year 1955. This is a result of great importance: 1955 is the year of the military coup in which President Juan Domingo Perón was overthrown by the military thus concluding a defining chapter in Argentine history. Further, we also find one structural break in legislative elections (it is dated 1949).[61]

---

59 As a measure of volatility we use the power transformed absolute growth $|y_t|^d$

60 Our data shows no guerilla warfare before 1948 and after 1977.

61 We also find no structural breaks for assassinations and cabinet changes. Further, we also find one structural break in cabinet size (it is dated 1946) while in anti-government demonstrations we find two breaks dated 1954 and 1972 (see graph 2.15 in the Appendix).

With arguably one exception (anti-government demonstrations in 1972, which were motivated by demands for the return of Perón from exile), all the structural breaks in our political instability series occur during Perón governments. Perón was elected president three times. His first term is from 1946 to 1952. He is re-elected in 1951, his second term starts in 1952 and ends abruptly in 1955. His third term is between 1973 (allowed to return from country-regionplaceSpain after 18-year exile) and 1974 (suffers fatal heart attack.) Although marked by severe economic problems, the second term (1951 to 1955) is more often remembered by the political instability (the various terrorist attacks being a sad prelude to the so-called “Dirty War” of 1970s.)

In what follows, we incorporate dummy variables in the equations (2.10), (2.11) (2.12) and (2.15), thus taking into account breaks in the political instability variables and in the volatility of growth. First, we introduce the following notation. D_{1t}, D_{2t} are (intercept) dummies defined as D_{1t}, $D_{2t} = 1$ in the periods 1922-2000 and 1964-2000, respectively, and D_{1t}, $D_{2t} = 0$ otherwise. Similarly, D_{it} is a (slope) dummy indicating the period which starts from the year of the break in the political instability variable (x_{it}). For example for strikes $D_{it} = 1$ in the period from 1955 to 2000 whereas for legislative elections $D_{it} = 1$ during the period from 1949 until the end of the sample.
The augmented model is given by

$$y_t = c + c_1 D_{1t} + c_2 D_{2t} + kh_t + \lambda x_{it} + \lambda_d D_{it} x_{it} + \epsilon_t, \tag{2.16}$$

and

$$h_t^{\frac{\delta}{2}} = \omega + \omega_1 D_{1t} + \omega_2 D_{2t} + \alpha h_{t-1}^{\frac{\delta}{2}} f\left(e_{t-1}\right) + \beta h_{t-1}^{\frac{\delta}{2}} + \phi x_{it} + \phi_d D_{it} x_{it} + \gamma y_{t-l}. \tag{2.17}$$

Recall that the coefficients φ and λ capture the impacts of the political instability variable on growth and its volatility respectively. Similarly, φ_d and λ_d correspond to the two effects from the year of the break onwards. Thus the two effects are captured by φ and λ in the period up to the year of the structural break, and by $\varphi + \varphi_d$ and $\lambda + \lambda_d$ during the period from the year of the break until the end of the sample. As above in order to study the direct effects of political instability and financial development we specify model 1 with $\varphi = \varphi_d = 0$, while model 2 with $\lambda = \lambda_d = 0$ allows us to investigate their indirect impacts on growth.

We also incorporate intercept dummies and level effects in the error correction equation (2.12) and conditional variance equation (2.15), as follows

$$\Delta y_t = \mu + \mu_1 D_{1t} + \mu_2 D_{2t} + \theta \Delta x_{i,t-l} + \varphi(y_{t-1} - c - \zeta x_{i,t-1}) + \varepsilon_t, \tag{2.18}$$

$$h_t^{\frac{\delta}{2}} = \omega + \omega_1 D_{1t} + \omega_2 D_{2t} + \alpha h_{t-1}^{\frac{\delta}{2}} \left| e_{t-1} \right|^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}} + \gamma y_{t-l}. \tag{2.19}$$

Overall, we find our results to be quite robust to the inclusion of the structural break dummies. That is, (i) informal instability (either guerilla warfare or strikes) and trade openness have a direct negative effect on growth, while formal instability (constitutional changes) have an indirect (through volatility) negative impact on growth (ii) trade openness affects growth negatively both directly and indirectly (see tables 2.19, and 2.20), (iii) the negative effects of the informal instability are significantly stronger in the short- than in the long-run, (iv) financial development affects growth positively in the long-run but negatively in the short-run (see table 2.21), (v) both the short- and long-run impact of the UK interest rate is negative, while trade openness does not affect growth in the long-run (see table 2.21).

It is also noteworthy that the causal negative effect of strikes reflects the period 1955-2000, which is not surprising given the intricate relationship between the Peron government and organized labor (see the λ_1 and λ_d columns in Table 2.39 in the Appendix of Campos et al., 2011). Interestingly, before 1949 there is no causal effect from legislative elections to growth volatility, whereas after 1949 a negative impact began to exist.

Finally, the only differences with the previous results is that the direct(indirect) effect of the UK interest rate(public deficit) disappears when we take into account structural breaks.

Table 2.19. Direct Effect of Guerrila Warfare/Strikes, Private Deposits/Savings Bank Deposits,UK Interest Rate and Trade Openness on Economic Growth. Dummies and (P)ARCH estimates

$x_{i,t}^{(fd)} \downarrow$	c_1	c_2	k	λ_1	λ_2	λ_3	λ_4	ω_1	ω_2	α	β	δ
							Guerrila Warfare					
Private Deposits/GDP	-0.032 (2.38)	0.054 (4.48)	0.689 (1.06)	-0.001 (4.01)	0.422 (3.17)	-0.0002 (0.39)	-0.028 (4.05)	–	0.031 (3.19)	0.71 (1.89)	0.36 (3.07)	1.00 –
Savings Bank Deposits/GDP	–	0.059 (6.99)	-1.060 (1.22)	-0.001 (3.52)	0.237 (3.87)	-0.0010 (1.08)	-0.026 (2.69)	–	0.038 (2.82)	0.73 (5.15)	0.38 (2.82)	0.80 –
							Strikes					
Private Deposits/GDP	–	0.063 (5.69)	1.121 (0.80)	-0.002 (2.67)	0.180 (0.44)	-0.0010 (0.36)	-0.055 (2.50)	–	0.017 (3.83)	0.41 (2.83)	0.60 (3.83)	1.10 –
Savings Bank Deposits/GDP	-0.009 (15.4)	–	0.893 (1.00)	-0.001 (1.76)	0.375 (2.52)	-0.0010 (1.17)	-0.048 (4.28)	0.015 (3.50)	0.074 (1.90)	0.84 (5.46)	0.08 (1.19)	1.00 –

Note: This table reports parameter estimates for the following model:

$y_t = c + c_1 D_{1t} + c_2 D_{2t} + kh_t + \lambda_1 x_{it}^{(pi)} + \lambda_2 x_{it}^{(fd)} + \lambda_3 x_{uk,t} + \lambda_4 x_{to,t} + \varepsilon_t,$

$h_t^{\frac{\delta}{2}} = \omega + \omega_1 D_{1t} + \omega_2 D_{2t} + \alpha h_{t-1}^{\frac{\delta}{2}} \mid e_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}}.$

$x_{it}^{(pi)}$ denotes an informal political instability (either guerilla warfare or strikes) variable.

$x_{it}^{(fd)}$ denotes a financial development (either private deposits/GDP or savings bank deposits/GDP) variable.

$x_{uk,t-1}$ is UK interest rate, and $x_{to,t-1}$ is trade openness.

D_{1t}, D_{2t} are (intercept) dummies defined as $D_{1t}, D_{2t} = 1$ in the periods 1922-2000 and 1964-2000 respectively, and $D_{1t}, D_{2t} = 0$ otherwise.

The numbers in parentheses are absolute t statistics.

Table 2.20. Indirect Effect of Constitutional Changes/Legislative Elections, UK Interest Rate, Trade Openness, and Public Deficit on Economic Growth. Dummies and (P)ARCH estimates

	c_1	c_2	k	ω_1	ω_2	α	β	ϕ_1	ϕ_2	ϕ_3	ϕ_4	γ	δ
Constitutional Changes	–	–	5.219 (3.67)	–	0.031 (3.37)	0.241 (1.90)	0.548 (4.30)	−0.003 (3.49)	0.005 (1.89)	−0.076 (3.07)	0.018 (1.36)	0.095 (3.43) $l=7$	1.00 –
Legislative Elections	−0.013 (18.59)	–	2.448 (1.24)	−0.002 (0.76)	0.044 (1.36)	0.261 (0.49)	0.558 (15.50)	−0.003 (5.02)	−0.001 (0.21)	−0.107 (2.46)	0.314 (2.75)	0.026 (0.30) $l=7$	1.00 –

Note: This table reports parameter estimates (interest) for the following model: $y_t = c + c_1 D_{1t} + c_2 D_{2t} + kh_t + \varepsilon_t$,

$h_t^{\frac{\delta}{2}} = \omega + \omega_1 D_{1t} + \omega_2 D_{2t} + \alpha h_{t-1}^{\frac{\delta}{2}} \mid e_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}} + \phi_1 x_{it}^{(pi)} + \phi_2 x_{uk,t} + \phi_3 x_{to,t} + \phi_4 x_{pd,t} + \gamma y_{t-l}$,

where $x_{it}^{(pi)}$ indicates a formal political instability variable (either constitutional changes or legislative elections),

$x_{uk,t}$ is UK interest rate, $x_{to,t}$ is trade openess, and $x_{pd,t}$ is public deficit.

D_{1t}, D_{2t} are (intercept) dummies defined as $D_{1t}, D_{2t} = 1$ in the periods 1922-2000 and 1964-2000 respectively, and $D_{1t}, D_{2t} = 0$ otherwise.

The numbers in parentheses are absolute t statistics.

Table 2.21. The short- and long-run Growth effects of Guerrila Warfare/Strikes, Private Deposits/ Savings Bank Deposits, UK Interest Rate, and Trade Openness with Dummies

$x_{it}^{(fd)} \downarrow$	μ_1	μ_2	θ_1	θ_2	θ_3	θ_4	φ	ζ_1	ζ_2	ζ_3	ζ_4	ω_1	ω_2	δ
							Guerrila Warfare							
Private Deposits/GDP	–	0.050 (4.67)	−0.0012 (7.73) $l=0$	−0.112 (4.30) $l=2$	−0.005 (19.57) $l=0$	−0.005 (2.22) $l=0$	−0.485 (9.96)	−0.0012 (3.65)	0.2348 (1.51)	−0.0034 (3.05)	−0.0084 (1.40)	−0.125 (2.64)	0.056 (4.26)	0.80 –
Savings Bank Deposits/GDP	–	0.050 (8.73)	−0.0003 (5.16) $l=3$	−0.236 (3.03) $l=3$	−0.003 (12.12) $l=0$	−0.023 (9.90) $l=0$	−0.481 (10.10)	−0.0009 (2.84)	0.1708 (2.21)	−0.0039 (5.54)	−0.0108 (1.43)	–	0.042 (3.08)	0.80 –
							Strikes							
Private Deposits/GDP	–	0.042 (5.13)	−0.0006 (3.06) $l=3$	−0.923 (6.11) $l=3$	−0.003 (3.82) $l=3$	−0.038 (6.86) $l=0$	−0.410 (8.80)	−0.0010 (2.23)	0.3205 (1.57)	−0.0039 (4.95)	−0.0059 (0.67)	–	0.033 (4.10)	0.90 –
Savings Bank Deposits/GDP	–	0.050 (3.37)	−0.0007 (4.51) $l=0$	−0.299 (4.38) $l=1$	−0.003 (5.17) $l=3$	−0.031 (4.89) $l=2$	−0.689 (8.78)	−0.0006 (1.95)	0.3779 (3.47)	−0.0038 (5.08)	−0.0092 (1.19)	–	0.047 (4.08)	1.00 –

Note: This table reports parameter estimates (of interest) for the following model:

$$\Delta y_t = \mu + \mu_1 D_{1t} + \mu_2 D_{2t} + \theta_1 \Delta x_{i,t-l}^{(pi)} + \theta_2 \Delta x_{i,t-l}^{(fd)} + \theta_3 \Delta x_{uk,t-l} + \theta_4 \Delta x_{to,t-l} + \varphi(y_{t-1} - c - \zeta_1 x_{i,t-1}^{(pi)} - \zeta_2 x_{i,t-1}^{(fd)} - \zeta_3 x_{uk,t-1} - \zeta_4 x_{to,t-1}) + \varepsilon_t,$$

$h_t^{\frac{\delta}{2}} = \omega + \omega_1 D_{1t} + \omega_2 D_{2t} + \alpha h_{t-1}^{\frac{\delta}{2}} |u_{t-1}|^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}}$. The four θ's (l is the order of the lag) and the four ζ's capture the short- and long-run effects respectively.

φ indicates the speed of adjustment to the long-run relationship.

$x_{i,t-l}^{(fd)}$ denotes a financial development (either private deposits/GDP or savings bank deposits/GDP) variable.

$x_{i,t-l}^{(pi)}$ denotes an informal political instability (either guerilla warfare or strikes) variable.

$x_{uk,t-l}$ is UK interest rate, $x_{to,t-l}$ is trade openness.

D_{1t}, D_{2t} are (intercept) dummies defined as $D_{1t}, D_{2t} = 1$ in the periods 1922-2000 and 1964-2000 respectively, and $D_{1t}, D_{2t} = 0$ otherwise.

The numbers in parentheses are absolute t statistics.

2.5.5 SUMMARY OF RESULTS

Out of the five informal political instabiliy variables used in Campos et al. (2011a) two, namely guerilla warfare and strikes, have powerful and precise negative effects on growth, which remain if we add to the baseline specification any of our additional variables. The negative growth effects of the informal instability are significantly stronger in the short- than in the long-run. The direct negative impact of strikes on growth reflects the 1955-2000 period.

From the six measures of formal political instability used in Campos et al. (2011a) only constitutional changes have a dominant indirect negative influence on growth.

From the four financial development variables employed in Campos et al. (2011a) the two measures of financial efficiency (private and saving bank deposits) rather than the two measures of the size of the financial sector are the primary influences. They affect growth positively in the long-run but negatively in the short-run.

From the four additional variables, namely inflation, public deficit, UK interest rate and trade openness, only the last two play an important role for growth. Both short- and long-run effects of the UK interest rate are negative whereas the indirect influence is positive. However, when we take into account structural breaks the negative direct impact disappears. Finally both direct and indirect effects of trade openness are negative but in the long-run the impact disappears.

Table 2.22. Summary of Results: Direct, Indirect, Short- and Long-run Effects.

↓Effects	Informal Instability (Guerilla Warfare/Strikes)	Formal Political Instability (Constitutional Changes)	Financial Efficiency (Private or Saving Bank Deposits)	UK Interest Rate	Trade Openess
Direct	Negative (For Strikes: It reflects the period 1955-2000)	-	Positive	Negative (Disappears with structural breaks)	Negative
Indirect	-	Negative	-	Positive	Negative
Short-run	Negative (For Strikes stronger after 1955)	-	Negative	Negative	Negative
Long-run	Negative (Smaller than the short-run)	-	Positive	Negative	Positive (Disappears in the multivariate analysis)

Note: The negative short- and long-run effects of public deficit disappears in the multivariate analysis.
Similarly the positive indirect impact of public deficit disappears when we take into account structural breaks.

2.6 Conclusions and Future Research

Using a PARCH framework and data for Argentina from approximately 1890 to 2000 we ask the following questions: What is the relationship between, on the one hand, financial development (domestic and international), inflation, public deficit, trade openness and political instability and, on the other hand, economic growth and (predicted) growth volatility? Are these effects fundamentally and systematically different? Does the intensity and the direction (the sign) of these effects vary over time, in general and, in particular, do they vary with respect to short- versus long-run considerations? We find that the main explanatory factors, solely in terms of their direct effects on economic growth in Argentina, turn out to be financial efficiency, informal political instability (either guerilla warfare or strikes), the UK interest rate and trade openness. However, the effect of the UK interest rate disappears once breaks are accounted for. Further, we find robust evidence that both formal political instability (constitutional changes), the UK interest rate and trade openness affect growth indirectly via its volatility. No other variables in our basic set exhibit such robust estimates of their indirect effects. From investigating whether dynamic considerations affect our conclusions, we find important differences in terms of short and long-run behavior of our key variables, more specifically, while the effects of informal political instability and the UK interest rate (negative) are similar in the long- and short-run, that of financial development and trade openness are negative in the short- and positive in the long-run. However, the long-run effect of trade openness disappears in the multivariate analysis.

These findings are interest in themselves but they also matter because they raise a number of new questions that we believe may be useful in motivating future research. Here we highlight two suggestions. Regarding the role of finance in the process of economic development, our finding reinforces a large body of previous research in that we also show a strong, positive impact of financial development on growth in the long-run. We find that different forms of political instability affect growth through different channels over different time windows, making up for a strong and rather resilient effect that seem really too powerful vis-à-vis the benefits brought to the table by financial development. We can not forget however that Argentina is unique: no other country in the world since the Industrial Revolution went from riches to rags. Put it differently, Argentina is an outlier and further research could try to replicate our analysis using the historical experience of other countries (ideally in a panel setting). That is, to study the relationship between financial development and economic growth in a panel of developing countries would strengthen what we know. Yet, the data requirements are very heavy indeed, with most developing countries lacking historical data even on key figures, such as per capita GDP, going back to the beginning or middle of the XIXth century. This, of course, does not make this task less important.

The second suggestion refers to a possible methodological improvement, namely the application of the bivariate GARCH model to the problem at hand (albeit the relatively small number of observations). The joint estimation of the political instability-financial development-growth system in a panel of countries would clearly represent progress and is something we feel future research should try to address.

2.A Appendix

A. DIRECT EFFECTS

I. BIVARIATE ANALYSIS.

1. Formal political Instability and financial development

As can be seen from the first two rows of Table 2.23 anti-government demos and assassinations have no impact on growth (the λ_1's are insignificant) whereas guerilla warfare/strikes affect growth negatively in all four/three cases (see the λ_1's in the third and fourth rows of Table 2.23). Moreover, three out of the four financial development variables have a significant (positive) impact on growth. Only, M3/GDP has no impact on growth.

Table 2.23. Direct Effect of Informal Political Instability and Financial Development on Economic Growth

Financial Development ($x_{it}^{(fd)}$):	Private Deposits/GDP			Savings Bank Deposits/GDP			M3/GDP			M1/GDP		
	λ_1	λ_2	δ	λ_1	λ_2	δ	λ_1	λ_2	δ	λ_1	λ_2	δ
↓Informal Political Instability ($x_{it}^{(pi)}$)												
Anti-Government Demos	−0.0008	0.83	1.00	−0.0001	0.52	0.80	−0.0005	0.14	0.80	−0.0006	0.40	0.80
	(1.29)	(10.55)	–	(0.41)	(4.09)	–	(0.98)	(0.73)	–	(1.35)	(4.87)	–
Assassinations	0.0004	0.95	0.90	−0.0002	1.00	0.90	−0.0005	0.30	0.90	0.0005	0.51	0.80
	(0.92)	(9.01)	–	(0.27)	(4.20)	–	(0.35)	(0.96)	–	(0.54)	(2.90)	–
Guerilla Warfare	−0.0010	0.47	0.80	−0.0009	0.26	0.80	−0.0013	−0.23	0.90	−0.0007	0.36	0.80
	(4.78)	(3.86)	–	(3.02)	(1.26)	–	(3.53)	(0.69)	–	(2.25)	(4.73)	–
Strikes	−0.0006	0.81	0.80	−0.0005	0.68	0.80	−0.0011	−0.21	0.80	−0.0009	0.41	0.80
	(1.73)	(6.62)	–	(2.04)	(2.89)	–	(1.65)	(1.53)	–	(1.11)	(3.01)	–

Note: This table reports parameter estimates of λ_1,λ_2 and δ for the following model:

$$y_t = c + kh_t + \lambda_1 x_{it}^{(pi)} + \lambda_2 x_{it}^{(fd)} + \varepsilon_t,\, h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \mid e_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}}$$

$x_{it}^{(pi)}$ and $x_{it}^{(fd)}$ denote an informal political instability and a financial development variable respectively.

The numbers in parentheses are absolute t statistics.

2. Either Guerilla Warfare/Strikes or Financial Efficiency and UK Interest Rate

In the regressions in Table 2.24 we include UK interest rate with either informal political instability (guerilla warfare or strikes) or financial efficiency as explanatory variables. In almost all cases both variables have a significant effect on growth (see the λ_1 column). UK interest rate has a significant effect in two out of the four cases (see the λ_2 column).

Table 2.24. Direct Effect of either Guerilla Warfare/Strikes or Financial Efficiency and UK Interest Rate on Economic Growth

	k	λ_1	λ_2	α	β	δ
Panel A. Informal Political Instability ($x_{it}^{(pi)}$)						
Guerilla Warfare	2.08	−0.001	−0.005	0.89	0.40	0.90
	(2.86)	(3.45)	(18.08)	(5.69)	(6.32)	–
Strikes	0.96	−0.001	−0.002	0.69	0.53	0.90
	(2.32)	(1.70)	(1.30)	(3.68)	(3.49)	–
Panel B. Financial Efficiency ($x_{it}^{(fd)}$)						
Private Deposits/GDP	0.63	1.01	0.0002	0.89	0.51	0.90
	(2.70)	(3.19)	(0.19)	(4.69)	(4.05)	–
Savings Bank Deposits/GDP	0.67	0.55	−0.0020	0.88	0.49	0.80
	(1.66)	(2.27)	(1.95)	(4.30)	(2.87)	–

Note: This table reports parameter estimates for the following model:

$y_t = c + kh_t + \lambda_1 x_{it}^{(j)} + \lambda_2 x_{uk,t} + \varepsilon_t$, $h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \mid e_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}}$,

where the superscript j indicates either an informal political instability (pi)

or a financial development/efficiency (fd) variable and $x_{uk,t}$ is UK interest rate.

The numbers in parentheses are absolute t statistics.

3. Either Guerilla Warfare/Strikes or Financial Efficiency and Trade Openness

In the regressions in Table 2.25 in almost all cases both variables (trade openness and either informal political instability or financial efficiency) have a significant effect on growth (see the λ_1 and λ_2 columns).

Table 2.25. Direct Effect of either Guerilla Warfare/Strikes or Financial Efficiency and Trade Openness on Economic Growth

	k	λ_1	λ_2	α	β	δ
Panel A. Informal Political Instability ($x_{it}^{(pi)}$)						
Guerilla Warfare	0.98 (3.15)	−0.001 (2.52)	−0.05 (3.26)	0.66 (4.93)	0.55 (5.40)	0.80 –
Strikes	0.96 (2.89)	−0.001 (1.78)	−0.04 (2.47)	0.69 (3.60)	0.54 (3.78)	0.90 –
Panel B. Financial Efficiency ($x_{it}^{(fd)}$)						
Private Deposits/GDP	2.08 (3.11)	0.58 (2.70)	−0.04 (1.35)	0.64 (5.49)	0.62 (8.24)	0.80 –
Savings Bank Deposits/GDP	0.78 (1.80)	0.54 (2.22)	−0.04 (3.96)	0.81 (4.07)	0.56 (5.25)	0.90 –

Note: This table reports parameter estimates for the following model:

$y_t = c + kh_t + \lambda_1 x_{it}^{(j)} + \lambda_2 x_{to,t} + \varepsilon_t,\, h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \mid e_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}},$

where the superscript j indicates either an informal political instability (pi) or a financial development/efficiency (fd) variable and $x_{to,t}$ is trade openness.

The numbers in parentheses are absolute t statistics.

4. Either Guerilla Warfare/Strikes or Financial Efficiency and Public Deficit

When we control for either informal political instability or financial efficiency the effect of public deficit on growth disappears in three out of the four cases (see the λ_1 and λ_2 columns).

Table 2.26. Direct Effect of either Guerilla Warfare/Strikes or Financial Efficiency and Public Deficit on Economic Growth

	k	λ_1	λ_2	α	β	δ
Panel A. Informal Political Instability ($x_{it}^{(pi)}$)						
Guerilla Warfare	1.06	−0.0009	−0.01	0.65	0.54	0.80
	(3.41)	(1.87)	(0.64)	(3.71)	(2.99)	–
Strikes	1.05	−0.0025	−0.02	0.66	0.54	0.90
	(12.27)	(4.51)	(0.70)	(3.33)	(3.12)	–
Panel B. Financial Efficiency ($x_{it}^{(fd)}$)						
Private Deposits/GDP	0.84	0.92	−0.01	0.53	0.70	0.80
	(2.03)	(4.31)	(0.47)	(4.84)	(11.16)	–
Savings Bank Deposits/GDP	0.94	0.62	−0.03	0.72	0.56	0.80
	(1.73)	(3.92)	(2.67)	(4.15)	(3.57)	–

Note: This table reports parameter estimates for the following model:

$y_t = c + kh_t + \lambda_1 x_{it}^{(j)} + \lambda_2 x_{pd,t} + \varepsilon_t,\ h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \mid e_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}},$

where the superscript j indicates either an informal political instability (pi)

or a financial development/efficiency (fd) variable and $x_{pd,t}$ is public deficit.

The numbers in parentheses are absolute t statistics.

5. Either Guerilla Warfare/Strikes or Financial Efficiency and Inflation

As can be seen in Table 2.27 in the bivariate analysis growth is independent of changes in inflation (see the λ_2 column).

Table 2.27. Direct Effect of either Guerilla Warfare/Strikes or Financial Efficiency and Inflation on Economic Growth

	k	λ_1	λ_2	α	β	δ
Panel A. Informal Political Instability ($x_{it}^{(pi)}$)						
Guerilla Warfare	0.80	-0.0012	4×10^{-7}	0.95	0.43	1.00
	(3.12)	(4.52)	(0.01)	(6.16)	(2.83)	–
Strikes	1.08	-0.0013	-1×10^{-5}	0.67	0.56	1.00
	(2.36)	(1.38)	(0.25)	(3.94)	(4.91)	–
Panel B. Financial Efficiency ($x_{it}^{(fd)}$)						
Private Deposits/GDP	1.52	0.99	4×10^{-5}	0.79	0.56	0.90
	(2.14)	(5.79)	(0.94)	(4.47)	(6.79)	–
Savings Bank Deposits/GDP	0.61	0.59	3×10^{-6}	0.78	0.55	0.80
	(2.18)	(2.99)	(0.11)	(4.64)	(4.25)	–

Note: This table reports parameter estimates for the following model:

$$y_t = c + kh_t + \lambda_1 x_{it}^{(j)} + \lambda_2 x_{\pi t} + \varepsilon_t,\ h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \mid e_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}},$$

where the superscript j indicates either an informal political instability (pi) or a financial development/efficiency (fd) variable and $x_{\pi t}$ is inflation.

The numbers in parentheses are absolute t statistics.

II TRIVARIATE ANALYSIS.

1. Guerilla Warfare/Strikes, Financial Efficiency and either UK Interest Rate or Trade Openness

Table 2.28 reports the results from the trivariate analysis. That is, in the growth equation we include as explanatory variables: i) informal political instability, ii) financial efficiency and iii) either UK interest rate or trade openness. In almost all cases the three variables have a significant effect on growth (see the λ_1, λ_2 and λ_3 columns).

Table 2.28. Direct Effect of Guerilla Warfare/Strikes, Financial Efficiency,and either UK Interest Rate or Trade Openness on Economic Growth

		λ_1	λ_2	λ_3	δ	λ_1	λ_2	λ_3	δ
	$x_{it}^{(fd)} \downarrow$	UK Interest Rate ($x_{uk,t}$)				Trade Openness ($x_{to,t}$)			
Guerilla Warfare	Private Deposits/GDP	-0.001 (2.79)	0.39 (3.30)	-0.003 (3.33)	1.00 –	-0.001 (8.32)	0.49 (9.70)	-0.02 (4.56)	0.80 –
	Savings Bank Deposits/GDP	-0.001 (2.54)	0.30 (2.13)	-0.004 (4.49)	1.20 –	-0.001 (2.58)	0.09 (0.68)	-0.04 (3.62)	0.80 –
		UK Interest Rate ($x_{uk,t}$)				Trade Openness ($x_{to,t}$)			
Strikes	Private Deposits/GDP	-0.001 (4.64)	-0.01 (0.07)	-0.003 (3.92)	0.90 –	-0.001 (1.88)	0.41 (2.28)	-0.04 (2.34)	0.90 –
	Savings Bank Deposits/GDP	-0.001 (2.61)	0.34 (1.97)	-0.004 (4.70)	1.00 –	-0.001 (1.80)	0.36 (3.69)	-0.05 (4.92)	0.80 –

Note: This table reports parameter estimates of interest for the following model:

$y_t = c + kh_t + \lambda_1 x_{it}^{(pi)} + \lambda_2 x_{it}^{(fd)} + \lambda_3 x_{nt} + \varepsilon_t,\ h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \mid e_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}}$

where $x_{it}^{(pi)}$ is either guerilla warfare or strikes, $x_{it}^{(fd)}$ is either private deposits/GDP or savings bank deposits/GDP, and x_{nt} indicates either UK interest rate ($x_{uk,t}$) or trade openness ($x_{to,t}$).

The numbers in parentheses are absolute t statistics.

2. Guerilla Warfare, Financial Efficiency and either Inflation or Public Deficit

As with the bivariate analysis in the trivariate one, public deficit and inflation have no effect on growth (see the λ_3 columns in Table 2.29).

Table 2.29. Direct Effect of Guerilla Warfare, Financial Efficiency and Inflation/Public Deficit on Economic Growth

$x_{it}^{(fd)}$:	Private Deposits/GDP				Saving Bank Deposits			
x_{it}	λ_1	λ_2	λ_3	δ	λ_1	λ_2	λ_3	δ
Inflation	-0.0012	0.67	4×10^{-5}	0.80	-0.0007	0.84	1×10^{-5}	0.90
	(5.05)	(7.83)	(1.44)	–	(1.25)	(7.09)	(0.31)	–
Public Deficit	-0.0011	0.48	0.01	1.00	-0.0007	0.79	0.02	0.90
	(4.60)	(2.94)	(1.40)	–	(0.98)	(4.99)	(0.86)	–

Note: This table reports parameter estimates of interest for the following model:

$$y_t = c + kh_t + \lambda_1 x_{gw,t} + \lambda_2 x_{it}^{(fd)} + \lambda_3 x_{it} + \varepsilon_t,\ h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \mid e_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}}$$

where $x_{gw,t}$ is guerilla warfare, $x_{it}^{(fd)}$ is a financial efficiency variable,

and x_{nt} indicates either inflation ($x_{\pi t}$) or public deficit ($x_{pd,t}$).

The numbers in parentheses are absolute t statistics.

B. INDIRECT EFFECTS.

I. BIVARIATE ANALYSIS.

1. Formal Political Instability and UK Interest Rate

As can be seen from Table 2.30 when we include in the variance of growth UK interest rate and one of the four alternative measures of formal political instability the effects of cabinet changes and cabinet size disappears (see the ϕ_1 column).

Table 2.30. Indirect Effect of Formal Political Instability and UK Interest Rate on Economic Growth

$x_{it}^{(pi)} \downarrow$	k	α	β	ϕ_1	ϕ_2	γ	δ
Cabinet Changes	3.81 (1.81)	0.16 1.19	0.91 (10.19)	0.0016 (0.69)	0.016 (3.09)	−0.024 (0.22)	1.00 –
Cabinet Size	2.47 (2.97)	0.27 (1.67)	0.77 (8.18)	0.0003 (0.25)	0.011 (5.10)	0.042 (0.76)	1.00 –
Constitutional Changes	2.04 (2.36)	0.32 (5.36)	0.72 (18.62)	−0.0016 (7.31)	0.012 (5.11)	0.138 (6.61)	1.00 –
Legislative Elections	2.01 (3.21)	0.38 (4.10)	0.68 (8.13)	−0.0060 (3.21)	0.010 (4.65)	0.006 (1.62)	1.00 –

Note: This table reports parameter estimates for the following model:

$$y_t = c + kh_t + \varepsilon_t,\ h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \mid e_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}} + \phi_1 x_{it}^{(pi)} + \phi_2 x_{uk,t} + \gamma y_{t-6},$$

where $x_{it}^{(pi)}$ indicates a formal political instability variable and $x_{uk,t}$ is UK interest rate.

The numbers in parentheses are absolute t statistics.

2. Formal Political Instability and Trade Openness

As can be seen from Table 2.31 when we include in the variance of growth trade openness and one of the four alternative measures of formal political instability the effects of cabinet size disappears (see the ϕ_1 column). In addition, the effect of trade openness is significant in three out of the four cases (see the ϕ_2 column).

Table 2.31. Indirect Effect of Formal Political Instability and Trade Openness on Economic Growth

$x_{it}^{(pi)} \downarrow$	k	α	β	ϕ_1	ϕ_2	γ	δ
Cabinet Changes	0.90 (2.43)	0.88 (6.32)	0.34 (4.66)	-0.0037 (5.16)	-0.08 (4.74)	0.180 (5.12)	1.00 –
Cabinet Size	1.22 (2.48)	0.75 (4.98)	0.42 (6.57)	-0.0001 (0.90)	-0.06 (4.14)	0.200 (4.64)	1.00 –
Constitutional Changes	2.50 (2.01)	0.48 (2.34)	0.53 (2.29)	-0.0142 (4.62)	-0.10 (1.20)	-0.002 (0.02)	0.90 –
Legislative Elections	2.93 (2.56)	0.26 (2.09)	0.78 (5.36)	-0.0100 (2.06)	-0.16 (2.53)	-0.008 (0.11)	0.90 –

Note: This table reports parameter estimates for the following model:

$$y_t = c + kh_t + \varepsilon_t,\ h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \mid e_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}} + \phi_1 x_{it}^{(pi)} + \phi_2 x_{to,t} + \gamma y_{t-6},$$

where $x_{it}^{(pi)}$ indicates a formal political instability variable and $x_{to,t}$ is trade openness.

The numbers in parentheses are absolute t statistics.

3. Formal Political Instability and Public Deficit

In the regressions in Table 2.32 we include formal political instability and public deficit as explanatory variables. Both variables have a significant effect on growth volatility (see the ϕ_1 and ϕ_2 columns).

Table 2.32. Indirect Effect of Formal Political Instability and Public Deficit on Economic Growth

$x_{it}^{(pi)} \downarrow$	k	α	β	ϕ_1	ϕ_2	γ	δ
Cabinet Changes	1.23	0.549	0.54	−0.0034	0.01	0.21	1.00
	(2.22)	(3.00)	(5.39)	(2.28)	(0.86)	(3.54)	–
Cabinet Size	1.02	0.674	0.50	−0.0002	0.06	0.19	1.00
	(2.07)	(3.23)	(4.88)	(4.52)	(4.90)	(4.05)	–
Constitutional Changes	1.57	0.613	0.48	−0.0066	0.08	0.09	1.00
	(2.22)	(3.15)	(3.06)	(1.73)	(3.72)	(1.19)	–
Legislative Elections	−0.44	0.001	1.03	−0.0094	0.01	0.04	0.80
	(1.55)	(0.02)	(25.49)	(1.84)	(2.21)	(0.57)	–

Note: This table reports parameter estimates for the following model:

$$y_t = c + kh_t + \varepsilon_t,\ h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \mid e_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}} + \phi_1 x_{it}^{(pi)} + \phi_2 x_{pd,t} + \gamma y_{t-6},$$

where $x_{it}^{(pi)}$ indicates a formal political instability variable and $x_{pd,t}$ is public deficit.

The numbers in parentheses are absolute t statistics.

4. Formal Political Instability and Inflation

When we control for formal political instability the effect of inflation on growth volatility disappears. That is, the ϕ_2 estimated coefficients are insignificant (See Table 2.33).

Table 2.33. Indirect Effect of Formal Political Instability and Inflation on Economic Growth

$x_{it}^{(pi)} \downarrow$	k	α	β	ϕ_1	ϕ_2	γ	δ
Cabinet Changes	1.70 (2.07)	0.44 (6.56)	0.55 (12.90)	-0.0064 (4.99)	5×10^{-5} (1.47)	0.19 (2.98)	1.00 –
Cabinet Size	1.96 (3.36)	0.61 (3.17)	0.56 (6.70)	-0.0003 (13.42)	6×10^{-5} (2.28)	0.17 (8.89)	1.00 –
Constitutional Changes	1.37 (1.54)	0.58 (5.33)	0.51 (5.81)	-0.0090 (3.82)	2×10^{-5} (0.58)	0.16 (2.06)	1.00 –
Legislative Elections	2.86 (2.63)	0.38 (4.99)	0.54 (8.45)	-0.0161 (2.37)	6×10^{-5} (1.96)	0.16 (5.70)	0.90 –

Note: This table reports parameter estimates for the following model:

$$y_t = c + kh_t + \varepsilon_t, \; h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \mid e_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}} + \phi_1 x_{it}^{(pi)} + \phi_2 x_{\pi t} + \gamma y_{t-6},$$

where $x_{it}^{(pi)}$ indicates a formal political instability variable and $x_{\pi t}$ is inflation.

The numbers in parentheses are absolute t statistics.

II. TRIVARIATE ANALYSIS.

Formal Political Instability, UK Interest rate and Trade Openness

As with the bivariate analysis when we control for UK interest rate and trade openness the effects of cabinet changes and cabinet size disappear (see table 2.34). That is, the estimates of ϕ_1 in the first two rows are insignificant.

Table 2.34. Indirect Effect of Formal Political Instability, UK Interest Rate and Trade Openness on Economic Growth

$x_{it}^{(pi)} \downarrow$	k	α	β	ϕ_1	ϕ_2	ϕ_3	γ	δ
Cabinet Changes	3.15 (2.59)	0.34 (2.22)	0.65 (7.45)	-0.0001 (0.02)	0.009 (4.76)	-0.08 (3.16)	0.24 (0.35) $l=7$	1.00 –
Cabinet Size	1.96 (1.77)	0.53 (3.44)	0.52 (3.23)	-0.0001 (0.62)	0.007 (1.91)	-0.07 (2.03)	0.01 (0.15) $l=5$	1.00 –
Constitutional Changes	1.62 (2.49)	0.36 (2.36)	0.77 (6.35)	-0.0112 (1.92)	0.012 (1.88)	-0.07 (4.35)	0.01 (0.12) $l=6$	1.00 –
Legislative Elections	3.00 (2.39)	0.32 (5.05)	0.71 (10.75)	-0.0056 (2.17)	0.012 (3.70)	-0.06 (2.04)	0.08 (0.98) $l=7$	1.00 –

Note: This table reports parameter estimates for the following model:

$y_t = c + kh_t + \varepsilon_t,$

$h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \mid e_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}} + \phi_1 x_{it}^{(pi)} + \phi_2 x_{uk,t} + \phi_3 x_{to,t} + \gamma y_{t-l},$

where $x_{it}^{(pi)}$ indicates a formal political instability variable,

$x_{uk,t}$ is UK interest rate and $x_{to,t}$ is trade openness.

The numbers in parentheses are absolute t statistics.

FOUR VARIABLES.

Constitutional Changes/Legislative Elections, UK Interest rate, Trade Openness and Inflation

When we include four explanatory variables in the variance of growth we find that inflation has no effect on growth (see the ϕ_4 column in Table 2.35). This result is consistent with the findings from the bivariate and trivariate analyses.

Table 2.35. Indirect Effect of Formal Political Instability, UK Interest Rate, Trade Openness, and Inflation on Economic Growth

	k	α	β	ϕ_1	ϕ_2	ϕ_3	ϕ_4	γ	δ
Constitutional Changes	3.32 (4.63)	0.57 (3.05)	0.5 (3.30)	-0.0004 (0.21)	0.006 (3.04)	-0.079 (2.37)	0.0002 (0.52)	0.018 (0.54) $_{l=6}$	1.00 –
Legislative Elections	4.17 (2.15)	0.23 (2.43)	0.73 (9.73)	-0.0036 (2.29)	0.008 (1.59)	-0.085 (1.52)	0.0005 (0.06)	-0.004 (0.52) $_{l=7}$	1.00 –

Note: This table reports parameter estimates for the following model:

$y_t = c + kh_t + \varepsilon_t$, $h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \mid e_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}} + \phi_1 x_{it}^{(pi)} + \phi_2 x_{uk,t} + \phi_3 x_{to,t} + \phi_4 x_{\pi t} + \gamma y_{t-l}$,

where $x_{it}^{(pi)}$ indicates a formal political instability variable,

$x_{to,t}$ is trade openness, $x_{uk,t}$ is UK interest rate, and $x_{\pi t}$ is inflation.

The numbers in parentheses are absolute t statistics.

C. SHORT- AND LONG-RUN EFFECTS.

I. UNIVARIATE ANALYSIS.

Either Informal Political Instability or Financial Development

In our univariate analysis in almost all cases the short- and long-run growth effects of informal political instability and of financial development are significant. Assassinations and M3/GDP have a weak long-run effect on growth (see columns θ and ζ in Table 2.36).

Table 2.36. The Short- and Long-run Growth Effects of Informal Political Instability/Financial Development

	θ	φ	ζ	α	β	δ
Panel A. Informal Political Instability ($x_{it}^{(pi)}$)						
Anti-Government Demos	-0.0036 (1.81) $l=5$	-0.43 (6.69)	-0.0017 (2.37)	0.84 (4.11)	0.51 (4.58)	0.80 –
Assassinations	-0.0010 (2.38) $l=2$	-0.64 (4.52)	-0.0007 (1.54)	1.18 (4.45)	0.34 (2.88)	0.90 –
Guerilla Warfare	-0.0011 (7.00) $l=0$	-0.74 (7.03)	-0.0017 (3.83)	1.14 (5.76)	0.34 (5.24)	0.90 –
Strikes	-0.0014 (4.01) $l=0$	-0.65 (6.34)	-0.0015 (3.26)	1.16 (5.81)	0.31 (4.08)	0.80 –
Panel B. Financial Development ($x_{it}^{(fd)}$)						
Private Deposits/GDP	-1.35 (1.81) $l=5$	-0.44 (4.64)	0.94 (23.72)	0.37 (2.63)	0.80 (6.69)	0.90 –
Savings Bank Deposits/GDP	-0.55 (1.89) $l=1$	-0.70 (3.23)	0.59 (4.84)	0.74 (6.69)	0.56 (6.21)	0.80 –
M3/GDP	-0.16 (3.00) $l=4$	-0.83 (4.11)	0.16 (1.60)	0.81 (6.59)	0.52 (7.19)	0.80 –
M1/GDP	-0.21 (1.91) $l=1$	-0.85 (4.14)	0.43 (4.20)	0.74 (6.89)	0.54 (6.62)	0.80 –

Note: This table reports parameter estimates for the following model:

$$\Delta y_t = \mu + \theta \Delta x_{i,t-l}^{(j)} + \varphi(y_{t-1} - c - \zeta x_{i,t-1}^{(j)}) + \varepsilon_t,$$

$h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \left|e_{t-1}\right|^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}}$. θ (l is the order of the lag) and ζ capture the short- and long-run effects respectively.

φ indicates the speed of adjustment to the long-run relationship.

The superscript j indicates either an informal political instability (pi) or a financial development (fd) variable.

The numbers in parentheses are absolute t statistics.

II. BIVARIATE ANALYSIS

Informal Political Instability and Financial Development

The results in Table 2.37 suggest that from the four informal political instability variables only guerilla warfare and strikes affect significantly growth in the long-run. That is when we control for the financial development the long-run effects of anti-government demos and assassinations disappear (see the ζ_2 columns). Similarly, M3/GDP has no long-run effect on growth in three out of the four cases (see the ζ_1 columns).

Table 2.37. The Short- and Long-run Growth Effects of Informal Political Instability and Financial Development

	θ_1	θ_2	φ	ζ_1	ζ_2	δ	θ_1	θ_2	φ	ζ_1	ζ_2	δ
	Guerrilla Warfare						Strikes					
Private Deposits/GDP	−0.26 (4.98) $l=1$	−0.0009 (6.87) $l=0$	−0.71 (5.78)	0.67 (6.39)	−0.0008 (2.99)	0.80 –	−0.67 (12.88) $l=1$	−0.0008 (8.63) $l=0$	−0.62 (6.76)	0.81 (6.66)	−0.0005 (1.74)	0.80 –
Savings Bank Deposits/GDP	−0.22 (1.62) $l=3$	−0.0015 (4.20) $l=3$	−0.63 (5.08)	0.29 (12.80)	−0.0006 (1.99)	0.90 –	−0.60 (2.00) $l=1$	−0.0021 (1.78) $l=3$	−0.50 (5.20)	0.60 (3.37)	−0.0002 (0.04)	0.80 –
M3/GDP	−0.25 (9.35) $l=0$	−0.0019 (4.54) $l=3$	−0.60 (9.11)	−0.07 (1.17)	−0.0013 (4.75)	0.90 –	−0.24 (1.99) $l=1$	−0.0014 (2.96) $l=0$	−0.70 (7.57)	0.15 (3.50)	−0.0010 (4.30)	0.90 –
M1/GDP	−0.24 (3.19) $l=1$	−0.0011 (2.85) $l=3$	−0.61 (7.29)	0.28 (6.00)	−0.0007 (3.21)	0.80 –	−0.22 (3.53) $l=0$	−0.0011 (2.98) $l=1$	−0.63 (7.41)	0.43 (3.64)	−0.0005 (1.68)	0.80 –
	Anti-Government Demos						Assassinations					
Private Deposits/GDP	−0.34 (9.01) $l=5$	−0.0036 (2.11) $l=5$	−0.57 (8.10)	0.82 (11.36)	−0.0007 (1.51)	0.90 –	−0.57 (2.00) $l=1$	−0.0008 (1.83) $l=1$	−0.70 (3.42)	0.93 (8.63)	0.0010 (1.43)	0.80 –
Savings Bank Deposits/GDP	−0.33 (1.97) $l=1$	−0.0011 (1.79) $l=1$	−0.64 (3.26)	0.51 (5.26)	0.0009 (1.61)	0.80 –	−0.89 (4.89) $l=1$	−0.0008 (1.88) $l=1$	−0.70 (4.05)	0.99 (4.06)	−0.0002 (0.18)	0.80 –
M3/GDP	−0.25 (2.70) $l=6$	−0.0009 (1.61) $l=1$	−0.62 (2.96)	0.04 (0.32)	0.0006 (0.29)	0.80 –	−0.14 (2.17) $l=1$	−0.0015 (1.57) $(l=4)$	−0.69 (2.58)	0.25 (1.55)	0.0007 (1.30)	0.80 –
M1/GDP	−0.33 (10.55) $l=1$	−0.0009 (2.07) $l=1$	−0.81 (4.00)	0.42 (6.30)	0.0004 (1.00)	0.80 –	−0.41 (3.02) $l=1$	−0.0007 (1.88) $l=1$	−0.70 (3.51)	0.47 (2.92)	0.0005 (1.03)	0.80 –

Note: This table reports parameter estimates for the following model:

$$\Delta y_t = \mu + \theta_1 \Delta x_{i,t-l}^{(fd)} + \theta_2 \Delta x_{i,t-l}^{(pi)} + \varphi(y_{t-1} - c - \zeta_1 x_{i,t-1}^{(fd)} - \zeta_2 x_{i,t-1}^{(pi)}) + \varepsilon_t,\ h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \left| e_{t-1} \right|^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}}$$

θ_1, θ_2 (l is the order of the lag) and ζ_1, ζ_2 capture the short-and long-run effects respectively.

φ indicates the speed of adjustment to the long-run relationship. $x_{i,t-l}^{(fd)}$ and $x_{i,t-l}^{(pi)}$ denote a financial development and an informal political instability variable respectively. The numbers in parentheses are absolute t statistics.

II. TRIVARIATE ANALYSIS.

Informal Political Instability, Financial Efficiency and either Trade Openness or UK Interest Rate

As with the univariate analysis, when we control for the informal political instability and financial efficiency, the short-run growth effects of UK interest rate and trade openness are still significant (see the θ_3 column in table 2.38) whereas in the long-run only the former affects growth significantly.(see the ζ_3 column)

Table 2.38. The Short- and Long-run Growth Effects of Guerrila Warfare/ Strikes, Financial Efficiency, and Trade Openness/UK Interest Rate

		θ_1	θ_2	θ_3	φ	ζ_1	ζ_2	ζ_3	δ
	$x_{it}^{(fd)} \downarrow$				Trade Openness ($x_{to,t}$)				
Guerrilla Warfare	Private Deposits/GDP	-0.001 (4.09) $l=3$	-0.25 (4.47) $l=3$	-0.044 (9.28) $l=0$	-0.68 (6.44)	-0.001 (4.01)	0.48 (4.03)	-0.011 (1.37)	0.80 –
	Savings Bank Deposits/GDP	-0.001 (4.20) $l=3$	-0.29 (2.10) $l=1$	-0.031 (8.11) $l=0$	-0.64 (5.26)	-0.001 (2.05)	0.31 (5.28)	-0.035 (4.87)	0.80 –
					UK Interest Rate ($x_{uk,t}$)				
	Private Deposits/GDP	-0.002 (3.03) $l=3$	-3.46 (2.57) $l=3$	-0.001 1.66) $l=2$	-0.60 (6.85)	-0.001 (2.73)	0.41 (4.65)	-0.003 (4.28)	0.80 –
	Savings Bank Deposits/GDP	-0.003 (1.87) $l=3$	-0.81 (2.90) $l=1$	-0.004 (5.75) $l=0$	-0.17 (3.85)	-0.001 (3.46)	0.15 (2.15)	-0.003 (4.81)	0.80 –
	$x_{i,t}^{(fd)} \downarrow$				Trade Openness ($x_{to,t}$)				
Strikes	Private Deposits/GDP	-0.001 (5.18) $l=0$	-0.12 (2.83) $l=2$	-0.050 (8.25) $l=0$	-7.20 (11.02)	-0.001 (2.36)	0.81 (5.14)	-0.007 (0.44)	0.80 –
	Savings Bank Deposits/GDP	-0.002 (2.16) $l=3$	-0.39 (2.04) $l=1$	-0.036 (2.20) $l=0$	-0.50 (5.08)	-0.001 (1.22)	0.38 (2.48)	-0.048 (0.87)	1.00 –
					UK Interest Rate ($x_{uk,t}$)				
	Private Deposits/GDP	-0.001 (8.81) $l=0$	-0.26 (6.33) $l=1$	-0.006 (25.78) $l=0$	-0.68 (16.54)	-0.001 (2.00)	0.40 (15.59)	-0.004 (4.94)	0.80 –
	Savings Bank Deposits/GDP	-0.001 (2.11) $l=0$	-0.49 (3.84) $l=1$	-0.006 (8.27) $l=0$	-0.31 (5.80)	-0.001 (1.83)	0.37 (1.83)	-0.004 (4.41)	0.80 –

Note: This table reports parameter estimates for the following model:

$$\Delta y_t = \mu + \theta_1 \Delta x_{i,t-l}^{(pi)} + \theta_2 \Delta x_{i,t-l}^{(fd)} + \theta_3 \Delta x_{n,t-l} + \varphi(y_{t-1} - c - \zeta_1 x_{i,t-1}^{(pi)} - \zeta_2 x_{i,t-1}^{(fd)} - \zeta_3 x_{n,t-1}) + \varepsilon_t,$$

$h_t^{\frac{\delta}{2}} = \omega + \alpha h_{t-1}^{\frac{\delta}{2}} \left| e_{t-1} \right|^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}}$. $\theta_1, \theta_2, \theta_3$ (l is the order of the lag) and $\zeta_1, \zeta_2, \zeta_3$ capture the short- and long-run effects respectively. φ indicates the speed of adjustment to the long-run relationship.

$x_{i,t-l}^{(pi)}$ is either guerilla warfare or strikes. $x_{i,t-l}^{(fd)}$ denotes a financial efficiency variable.

$x_{n,t-l}$ indicates either trade openness ($x_{to,t-l}$) or UK interest rate ($x_{uk,t-l}$). The numbers in parentheses are absolute t statistics.

D. STRUCTURAL BREAKS.

I. DIRECT EFFECT

Adding a dummy to take into account the break in strikes. The causal negative effect of strikes reflects the period 1955-2000, which is not surprising given the intricate relationship between the Peron government and organized labor (see the λ_1 and λ_d columns in Table 2.39). The other effects are robust to the inclusion of a slope dummy for strikes.

Table 2.39. Direct Effect of Strikes, Financial Efficiency, Trade Openness and UK Interest Rate on Economic Growth (with Dummy for Strikes)

	c_1	c_2	k	λ_1	λ_d	λ_2	λ_3	λ_4	ω_1	ω_2	α	β	δ
Private Deposits/GDP	–	–	1.85 (1.62)	−0.0003 (0.07)	−0.002 (2.86)	0.51 (2.70)	0.001 (1.01)	−0.008 (1.84)	−0.05 (2.74)	0.07 (3.21)	0.50 (4.09)	0.34 (2.95)	0.80 –
Savings Bank Deposits/GDP	−0.02 (4.68)	0.06 (4.96)	−1.15 (1.01)	−0.0003 (0.11)	−0.001 (1.77)	0.20 (1.73)	−0.003 (3.01)	−0.002 (0.26)	–	0.03 (2.66)	0.84 (5.01)	0.32 (3.96)	1.00 –

Note: This table reports parameter estimates for the following model:

$$y_t = c + c_1 D_{1t} + c_2 D_{2t} + k h_t + (\lambda_1 + \lambda_d D_{gs,t}) x_{gs,t} + \lambda_2 x_{it}^{(fd)} + \lambda_3 x_{uk,t} + \lambda_4 x_{to,t} + \varepsilon_t,$$

$$h_t^{\frac{\delta}{2}} = \omega + \omega_1 D_{1t} + \omega_2 D_{2t} + \alpha h_{t-1}^{\frac{\delta}{2}} \mid e_{t-1} \mid^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}}.$$

$x_{gs,t}$ denotes strikes. $x_{it}^{(fd)}$ denotes a financial development/efficiency variable

$x_{uk,t}$ is UK interest rate, and $x_{to,t}$ is trade openness.

D_{1t}, D_{2t} are (intercept) dummies defined as $D_{1t}, D_{2t} = 1$ in the periods 1922-2000 and 1964-2000 respectively, and $D_{1t}, D_{2t} = 0$ otherwise.

$D_{gs,t}$ is a slope dummy defined as $D_{gs,t} = 1$ in the period 1955-2000, and $D_{gs,t} = 0$ otherwise.

The numbers in parentheses are absolute t statistics.

SHORT- AND LONG-RUN EFFECTS.

The short- and long-run effects are robust to the inclusion of a slope dummy for strikes (see Table 2.40).

Table 2.40. The Short- and Long-run Growth effects of Strikes, Financial Efficiency, UK Interest Rate and Trade Openness (with Dummy for Strikes)

	μ_1	μ_2	θ_1	θ_d	θ_2	θ_3	θ_4	φ	ζ_1	ζ_2	ζ_3	ζ_4	ω_1	ω_2	δ
Private Deposits/GDP	–	0.05 (5.24)	−0.0004 (3.38) $l=1$	−0.001 (9.20)	−0.67 (4.99) $l=3$	−0.003 (5.50) $l=3$	−0.041 (6.86) $l=0$	−0.38 (6.25)	−0.0010 (2.23)	0.32 (1.57)	−0.004 (4.95)	−0.006 (0.67)	–	0.03 (5.04)	0.80 –
Savings Bank Deposits/GDP	–	0.05 (4.99)	−0.0004 (2.56) $l=0$	−0.001 (3.46)	−0.34 (3.18) $l=1$	−0.005 (8.77) $l=0$	−0.006 (1.90) $l=1$	−0.36 (5.62)	−0.0006 (1.95)	0.38 (3.47)	−0.004 (5.08)	−0.009 (1.19)	–	0.02 (3.99)	1.00 –

Note: This table reports parameter estimates (of interest) for the following model:

$\Delta y_t = \mu + \mu_1 D_{1t} + \mu_2 D_{2t} + (\theta_1 + \theta_d D_{gs,t-l})\Delta x_{gs,t-l} + \theta_2 \Delta x^{(fd)}_{i,t-l} + \theta_3 \Delta x_{uk,t-l} + \theta_4 \Delta x_{to,t-l} +$
$\varphi(y_{t-1} - c - \zeta_1 x_{gs,t-1} - \zeta_2 x^{(fd)}_{i,t-1} - \zeta_3 x_{uk,t-1} - \zeta_4 x_{to,t-1}) + \varepsilon_t,$
$h_t^{\frac{\delta}{2}} = \omega + \omega_1 D_{1t} + \omega_2 D_{2t} + \alpha h_{t-1}^{\frac{\delta}{2}} \left|e_{t-1}\right|^{\delta} + \beta h_{t-1}^{\frac{\delta}{2}}$. $\theta_1, \theta_2, \theta_3$ (l is the order of the lag) and $\zeta_1, \zeta_2, \zeta_3$ capture the short- and long-run effects respectively. φ indicates the speed of adjustment to the long-run relationship.

$x_{gs,t-l}$ is strikes. $x^{(fd)}_{i,t-l}$ denotes a financial development/efficiency variable.

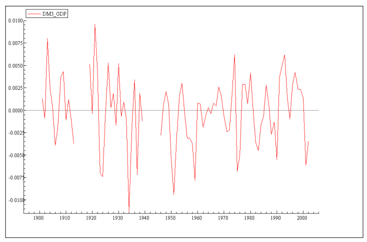
$x_{uk,t-l}$ is UK interest rate, $x_{to,t-l}$ is trade openness.

D_{1t}, D_{2t} are (intercept) dummies defined as $D_{1t}, D_{2t} = 1$ in the periods 1922-2000 and 1964-2000 respectively, and $D_{1t}, D_{2t} = 0$ otherwise.

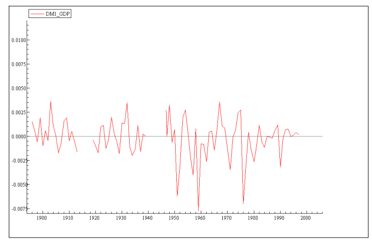
$D_{gs,t}$ is a slope dummies defined as $D_{gs,t} = 1$ in the period 1955-2000, and $D_{gs,t} = 0$ otherwise.

The numbers in parentheses are absolute t statistics.

Fig. 2.13. Measures of Financial Development

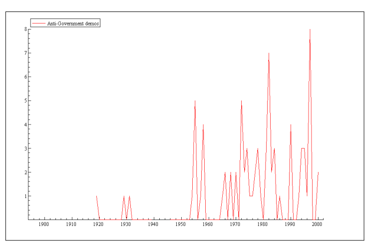


M3 over GDP

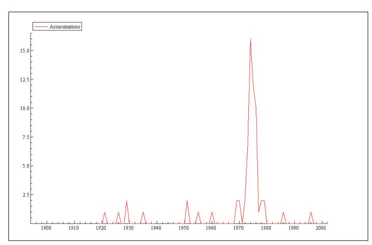


M1 over GDP

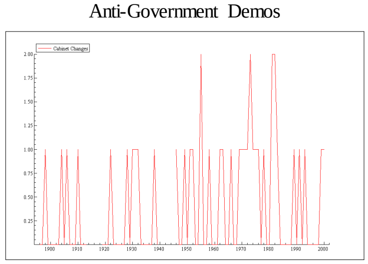
Fig. 2.14. Measures of Political Instability



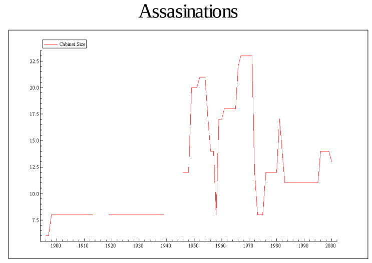
Anti-Government Demos



Assasinations

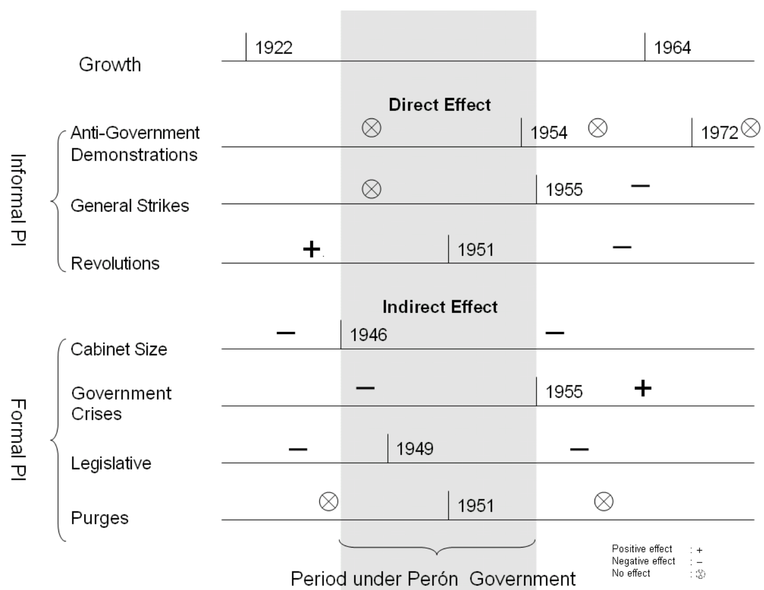


Cabinet Changes



Cabinet Size

Fig. 2.15. Breaks



Chapter 3
Are Economic Growth and the Variability of the Business Cycle Related? Evidence From Asia

3.1 Introduction

[62]Until the early 1980s, macroeconomic theorists treated the analysis of the real business cycle (RBC) as separate from the study of economic growth. In the 1980s, three important contributions in business cycle theory by Kydland and Prescott (1982), Long and Plosser (1983), and King et al. (1988) integrated the theories of the business cycle and economic growth in their models. However, these models did not consider the possibility that the variability of the business cycle might relate to the rate of economic growth. Similarly, for the most part, developments in growth theory have been made without consideration of the variability in the business cycle. The scene has changed recently at both the theoretical and empirical front. At the theoretical level, Blackburn and Pelloni (2005) and a number of studies summarised by these authors examine how the cyclical fluctuations might relate to long-run economic growth. At the empirical level, highlight the importance of the reduction in US GDP growth volatility in the last two decades and its implications for growth theory. The early dichotomy in macroeconomic theory between economic growth and the variability of economic fluctuations should be reconsidered given several theories outlined below regarding the relationship between output volatility and growth. These theories predict a positive, negative or no association between the two variables. The empirical evidence to date, based on cross-section country studies, panel

[62] This chapter is a sister paper of Fountas and Karanasos (2008) “Are economic growth and the variability of the business cycle related? Evidence from five European countries”, Introduction and literature sections are generally based on Fountas and Karanasos (2008).

data studies, or time-series analyses of individual countries is also quite mixed. The theoretical and empirical ambiguity surrounding the RBC variability–economic growth relationship provides us with the motivation to expand on the empirical aspects of this issue.

We attempt to cover a gap in the existing empirical literature by employing a monthly output data that starts in the 1960s on four Asian countries/regions. This approach allow us to analyze the RBC variability–growth relationship over a period that including significant variation in output growth, such as the 1970s oil shock (supply shock), the 2001 electronic industry depression (demand shock) and the credit crunch in 2008 (financial crisis). The rest of the chapter is structured as follows. Section 2 presents a survey of the theoretical literature on the relationship between the RBC and economic growth. Section 3 reviews the existing empirical literature. Section 4 outlines our econometric model and section 5 presents our main results and an interpretation. Section 6 describes some robustness tests. Finally, Section 7 summarizes our main conclusions.

3.2 Theoretical Background

Given the absence of a theoretical consensus, the anticipated relationship between output variability and economic growth remains an empirical issue. Macroeconomic theory offers three possible scenarios regarding the impact of the former on the latter. First, there is the possibility of independence between output variability and growth. In other words, the determinants of the two variables are different from each other. According to some business cycle models, output fluctuations around the natural rate are due to price misperceptions in response to monetary shocks. On the other hand, changes in the growth rate of output arise from real factors such as technology (Friedman 1968). The scenario of a negative association between output variability and average growth goes back to Keynes (1936) who argued that entrepreneurs, when estimating the return on their investment, take into consideration the fluctuations in economic activity. The larger the output fluctuations, the higher the perceived riskiness of investment projects and, hence, the lower the demand for investment and output growth. A similar result is obtained by the literature on sunspot equilibria (Woodford 1990). According to Bernanke (1983) and Pindyck (1991), the negative relationship between output volatility and growth arises from investment irreversibilities at the firm level. Ramey and Ramey (1991) show that in the presence of commitment to technology in advance, higher output volatility can lead to suboptimal ex post output levels by firms (due to uncertainty-induced planning errors) and hence, lower mean output and growth. Finally, the positive impact of output variability on growth can be justified by several economic theories. First, more income variability (uncertainty) would lead to a higher savings rate (Sandmo, 1970) for precautionary reasons, and hence, according to neoclassical growth theory, a higher equilibrium rate of economic growth. This argument has been advanced by Mirman (1971). An alternative explanation is due to Black (1987) and is based on the hypothesis that investments in riskier technologies will be pursued only if the expected return on these investments (average rate of output growth) is large enough to compensate for the extra risk. As real investment takes time to materialize. More recently, Blackburn (1999) using a model of endogenous growth generated by learning-by-doing shows that business cycle volatility raises the long-run growth of the economy.

The effect of output volatility on growth is not always unambiguous. A number of studies (Smith 1996; Grinols and Turnovsky 1998; Turnovsky 2000) show that, with preferences represented by a constant elasticity utility function, the growth rate is positively related to volatility provided the coefficient of risk aversion exceeds one. Smith (1996) shows that the sign of the growth-volatility relationship depends on whether the intertemporal elasticity of substitution exceeds or falls short of one. The above papers all refer to a closed economy. Turnovsky and Chattopadhyay (2003), in a stochastic general equilibrium small-open economy model of a developing country, examine the effect of output volatility on growth allowing for three additional types of variability (in the terms of trade, government spending and money supply) to have an impact on output growth. The theoretical model implies that output volatility has an ambiguous effect on growth. This result is confirmed by numerical simulations that show that the effect is small. The opposite type of causality, running from growth to output uncertainty, may also be examined in the present analysis. From a theoretical point of view, the sign of this causality relationship is negative. An increase in growth leads to more inflation (the short-run Phillips curve effect). Empirical evidence by Briault (1995) supports this effect. Furthermore, a higher inflation rate lowers output growth uncertainty (Ball et al. 1988). In the tradition of New Keynesian Economics, a higher inflation rate leads to more frequent optimal price adjustments and therefore nominal shocks cause smaller real effects as well as a lower output variability. Recently, a growing theoretical literature has developed that examines the correlation between average output growth and its variability in an endogenous growth set-up (Blackburn and Galinder 2003; Blackburn and Pelloni 2004, 2005). Blackburn and Galinder (2003) focus on the importance of the source of technological change for the sign of correlations between growth and its volatility. In a stochastic real growth model the authors show that positive (negative) correlation will most likely arise in a framework of internal (external) learning where the agents improve their productive efficiency by investing time in learning (benefit from knowledge spillovers taking place among agents). In a stochastic monetary growth model Blackburn and Pelloni (2004) show that the correlation between growth and its variability is a function of the type of shocks buffeting the economy. The study concludes that the correlation will be positive (negative) depending on whether the real (nominal) shocks dominate. In a richer setting, Blackburn and Pelloni (2005) use a stochastic monetary growth model with three different types of shocks (technology, preference and monetary) that have permanent effects on output due to wage contracts and endogenous technology. The authors show that output growth and output variability are negatively correlated irrespective of the type of shocks causing fluctuations in the economy.

3.3 Empirical Evidence

The empirical evidence to date on the association between output growth and its variability is mixed. Early studies employ cross-section (Kormendi and Meguire 1985) or pooled data (Grier and Tullock 1989) and find evidence for a positive association. Ramey and Ramey (1995) use a panel of 92 countries and a sample of OECD countries (for the 1960–1985 period) and find strong evidence that countries with higher output variability have lower growth. A similar result was obtained by Zarnowitz and Moore (1986) who divid the 1903–1981 period into six subperiods and compare high and low growth periods in terms of output variability (measured by the standard deviation of the annual growth rate in real GNP). In a recent study, Kneller and Young (2001), using a panel-data framework, find that output variability reduces growth. Turnovsky and Chattopadhyay (2003) find a similar, although small, effect in a sample of 61 developing countries allowing in their model for various types of volatility to have an impact on growth. More recent studies use the time series techniques of Generalized Autoregressive Conditional Heteroskedasticity (GARCH) models to proxy for output uncertainty rather than variability (Caporale and McKiernan 1996, 1998; Speight 1999). The first two papers use UK and US data, respectively, and find a positive association between output variability and growth, whereas the last paper uses UK data and finds no association. Grier and Perry (2000) using the GARCH-M model and monthly US data find no evidence that uncertainty about output affects the rate of growth. Henry and Olekalns (2002) find evidence in favour of a negative association using post-war real GDP data for the United States. Allowing for asymmetries, Grier et al. (2004) find US evidence for a positive effect. Fountas et al. (2002) find no evidence for an effect of output uncertainty on growth using data from Japan and a bivariate GARCH model that includes inflation and growth. This result is confirmed in a recent study by Fountas et al. (2004) using Japanese data and three different univariate GARCH models. The motivation for our empirical study comes from several factors. First, the inconclusiveness of the existing empirical time series literature, second the sparsity of evidence on the effect of growth on its uncertainty.

We, therefore, attempt to provide more robust evidence on the bi-directional relationship between the two variables using monthly data for four Asian countries/regions.

3.4 PARCH Model

We adopt PARCH estimation for our output data. Since its introduction by Ding et al. (1993), the PARCH model has been frequently applied. For example, Hentschel (1995) defined a parametric family of asymmetric GARCH formulations that nests the EGARCH and PARCH models. He and Teräsvirta (1999) considered a family of first-order asymmetric GARCH processes which includes the asymmetric PARCH (A-PARCH) as a special case. Brooks et al. (2000) analyzed the applicability of the PARCH models to national stock market returns for ten countries[63]. Laurent (2004) derives analytical expressions for the score of the A-PARCH model. The use of the PARCH model is now widespread in the literature (see, for example, Mittnik and Paolella, 2000, Giot and Laurent, 2003, Karanasos and Schurer, 2005, Karanasos and Kim, 2006, and Conrad et al., 2006, 2007).

Let y follow an autoregressive (AR) process augmented by a 'risk premium' defined in terms of volatility

$$\Phi(L)y_t = \phi_0 + kg(h_t) + \varepsilon_t \tag{3.20}$$

with

$$\varepsilon_t = e_t h_t^{\frac{1}{2}}$$

[63] It is worth noting that Fornari and Mele (1997) show the usefulness of the PARCH scheme in approximating models developed in continuous time as systems of stochastic differential equations. This feature of GARCH schemes has usually been overshadowed by their well-known role as simple econometric tools providing reliable estimates of unobserved conditional variances (Fornari and Mele, 2001).

where by assumption the finite order polynomial $\Phi(L) = \sum_{i=1}^{p} \phi L^i$has zeros outside the unit circle. In addition, $\{e_t\}$ are independent and identically distributed (i.i.d) random variables with $E\left(e_t\right) = E\left(e_t^2 - 1\right) = 0$. The conditional variance of output growth $\{y_t\}$, is positive with probability one and is a measurable function of the sigma-algebra $\sum_{t-1}$, which is generated by $\{y_{t-1}, y_{t-2}, ...\}$.

Furthermore, we need to choose the form in which the time-varying variance enters the specification of the mean to determine the "risk premium". This is a matter of empirical evidence. In the empirical results that follow we employ three specifications for the functional form of the 'risk premium' ($g\left(h_t\right) = h_t$, $g\left(h_t\right) = \sqrt{h_t}$, or $g\left(h_t\right) = ln\left(h_t\right)$).

Moreover, h_t is specified as an A-PARCH(1, 1) process with lagged output growth included in the variance equation

$$h_t^{\frac{\delta}{2}} = \omega + \alpha f\left(e_{t-1}\right) + \beta h_{t-1}^{\frac{\delta}{2}} + \gamma_l \pi_{t-1} \tag{3.21}$$

with

$$f\left(e_{t-1}\right) \equiv \left[\left|e_{t-1}\right| - \zeta e_{t-1}\right]^{\delta}$$

Where δ with $\delta > 0$ is the "heteroscedasticity" parameter, α and β are the ARCH and GARCH coefficients respectively, ζ with $|\zeta| < 1$ is the "leverage" term and γ_l is the "level" term for the lag of output growth. The model imposes a Box-Cox power transformation of the conditional standard deviation process and the asymmetric absolute residuals. The expected value of is given by:

$$E\left[f\left(e_{t-1}\right)\right] = \left\{ \begin{array}{l} \frac{1}{\sqrt{\pi}} \left[(1-\zeta)^{\delta} + (1+\zeta)^{\delta}\right] 2^{\left(\frac{\delta}{2}-1\right)} \Gamma\left(\frac{\delta+1}{2}\right), \text{if } e_{t-1} \overset{(i.d)}{\sim} N(0,1) \\ \frac{(r-2)^{\frac{\delta}{2}} \Gamma\left(\frac{r-\delta}{2}\right) \Gamma\left(\frac{\delta+1}{2}\right)}{\Gamma\left(\frac{r}{2}\right) 2\sqrt{\pi}} \left[(1-\zeta)^{\delta} + (1+\zeta)^{\delta}\right], \text{if } e_{t-1} \overset{(i.d)}{\sim} t_r(0,1) \end{array} \right\}$$

where N and t denote the Normal and student-t distributions respectively, r are the degrees of freedom of the student-t distribution and $\Gamma(.)$ is the Gamma function. The δth moment of the conditional variance is a function of the above expression (see Karanasos and Kim, 2006).

Within the A-PARCH model, by specifying permissible values for δ, α, β, ζ and γ_l in Equation (3.21), it is possible to nest a number of the more standard ARCH and GARCH specifications (see Ding et al., 1993, Hentschel, 1995, and Brooks et al., 2000). For example, in Equation (3.21) let $\delta = 2$ and $\zeta = \gamma_l = 0$ to get the GARCH model. In order to distinguish the general model in Equations (3.20) and (3.21) from a version in which $k = \gamma_l = \zeta = \beta = 0$, we will hereafter refer to the former as A-PARCH-in-mean-level (A-PARCH-ML) and the latter as PARCH.

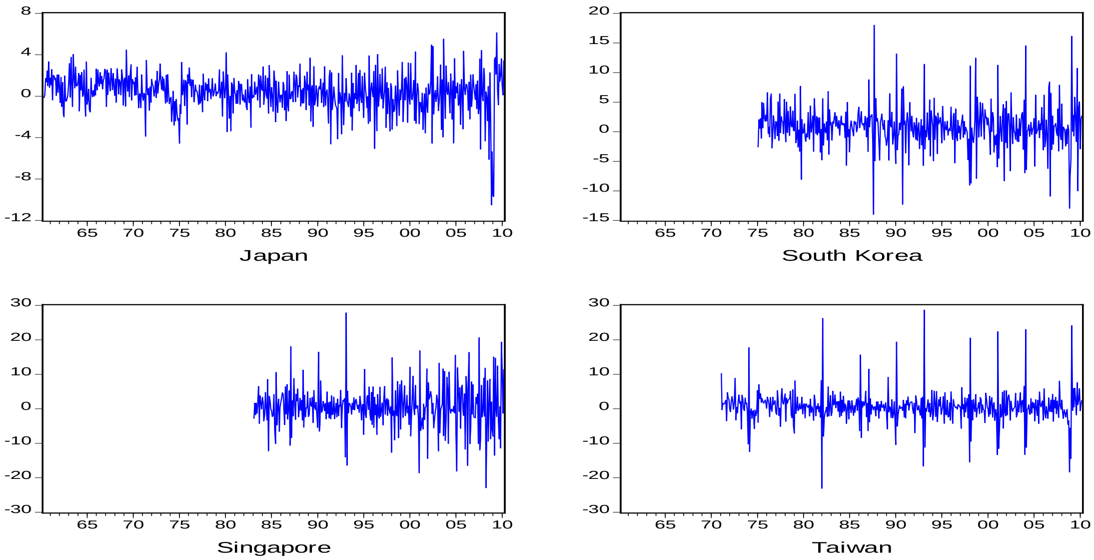
3.5 Empirical Analysis

3.5.1 Data and Power-transformed Growth

We use monthly data on the index of industrial production (IP) as proxies for the output and cover four Asian countries/regions, namely, Japan, South Korea, Singapore and Taiwan. The data end at March 2010, the starting point differs across countries/regions of our sample[64]. The source of our data series is Datastream. The growth rate of output is measured by the year-to-year changes in the log of Census X12 seasonal adjusted industrial production. Figure 3.16 plots the growth of the IP series in the four countries/regions. The choice of IP as a proxy for output is dictated by data availability considerations. It should be borne in mind that this proxy is not perfect since IP is about one quarter of real GDP and is more variable than the latter.

[64] April 1960, January 1975, January 1983 and January 1971 for Japan, South Korea, Singapore and Taiwan respectively.

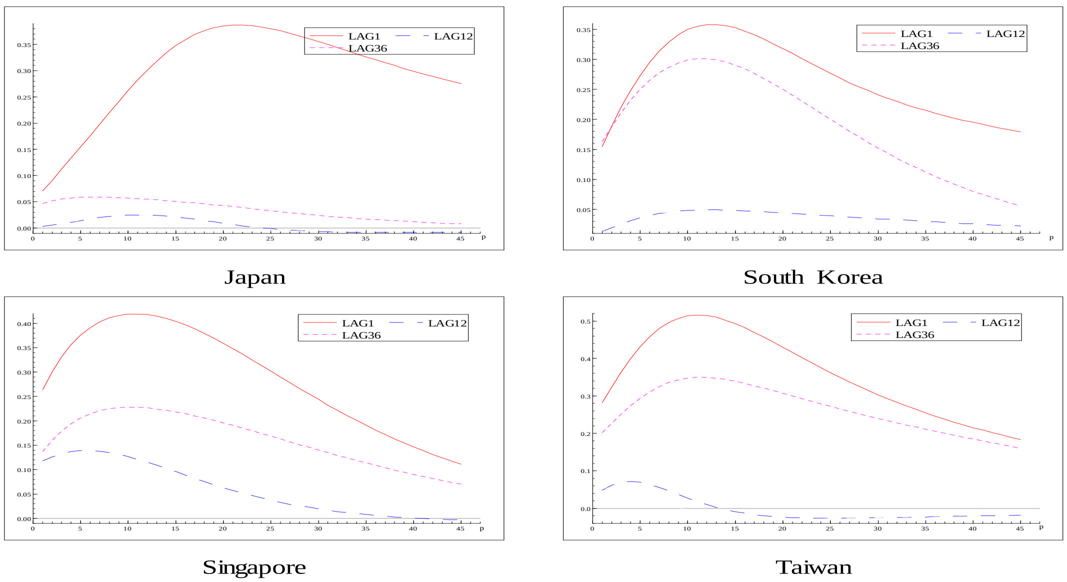
Fig. 3.16. Output Growth over Time



We employ the Augmented Dickey–Fuller (ADF) and Phillips–Perron (PP) unit root tests to test for the stationarity of all growth series as the estimation of GARCH models requires that all variables are stationary. Both unit root test results indicate that the four output growth series are I(0). We calculate the sample autocorrelations of the absolute value of growth $\rho_\tau(d)$ as a function of d for lags $\tau = 1, 12$ and 36, and taking $d = 0.1, 0.2, 0.3...$ to 4.5. Figure 3.17 provides the plots of the calculated $\rho_\tau(d)$. For example, for lag 36, there is a unique point d^* equal to $0.7, 1.2, 1.0$ and 1.4 for Japan, South Korea, Singapore and Taiwan respectively. Such that $\rho_{36}(d)$ reaches its maximum at this point: for $\rho_{36}(d^*) > \rho_{36}(d)$ for $d^* \neq d$. These figures indicate that in our data the autocorrelation structure of growth is the strongest for values of d smaller than two.[65].

[65] the lag1 of Japan data has maximum autocorrelation at $d = 2.1$.

Fig. 3.17. Autocorrelation of $|\pi_t|^d$ at Lag 1, 12, 36



3.5.2 Estimated Models of Growth

We proceed with the estimation of the AR-PARCH(1,1) model in equations (3.20) and (3.21) in order to take into account the serial correlation observed in the levels and power transformations of our time series data. The estimated parameters are obtained by quasi-maximum likelihood estimation (QMLE) as implemented in EVIEWS. The best fitting specification is chosen according to the Likelihood Ratio (LR) results and the minimum value of the Information Criteria (IC) (not reported). Once heteroscedasticity in the conditional mean is accounted for, an AR(12) specification appears to capture the serial correlation in all four growth series. Table 3.41 reports AR lags for the conditional means of inflation series.

The estimated standard errors are determined by the Bollerslev and Wooldridge (1992) procedure that accounts for the non-normality of the residuals. To check the sensitivity of our results to the form in which the time varying variance enters the specification of

Table 3.41. AR Lags

Japan	$1, 2, 3, 9, 10, 11$
South Korea	$1, 2, 7, 9, 12$
Singapore	$1, 2, 12$
Taiwan	$1, 3, 12$

Notes: The numbers represent the AR lags used in the mean equations of the model.

the mean, we not only use the conventional conditional variance, but also use either the conditional standard deviation or the logarithm of the conditional variance as regressor in the mean. In addition, we allow for the impact of inflation on growth and output uncertainty. Moreover, we allow for the possibility of structure break in the growth data. The mean equation and variance equation are modified to include structure break dummy variables. In other words, the dummy variables are included to reflect the difference between the low/negative periods (oil shock, demand shock, financial crisis) and the normal growth periods in the series. We find that dummies in all countries/regions are statistically significant in the mean equation but dummies are insignificant in the variance equation for South Korea, Singapore and Taiwan except Japan.

For all countries/regions, we find the leverage term ζ to be insignificant and therefore we re-estimate the model excluding this parameter. The estimated α parameter is highly significant in all cases. However, the β parameter is only significant in very small number of cases, so we re-estimate the model with ARCH term only. In order to distinguish the general PARCH model from a version in which δ is fixed to a specific value we will hereafter refer to the latter as (P)ARCH.

Table 3.42 reports the (P)ARCH models with or without ”level effect”. The IC chooses a (P)ARCH model with estimated power term parameter $\delta = 0.7, 0.5, 0.5$ and 1.1 of for Japan, South Korea, Singapore and Taiwan, respectively. The corresponding values for the model with lagged growth included in the conditional variance as the “level effect” (see “level” panel of Table 3.42) are lower for Japan: $\delta = 0.5$, but remarkably higher for Singapore, which the Akaike Information Criteria (AIC) choose (P)ARCH models with

Table 3.42. (P)ARCH Models

	Japan	South Korea		Singapore	Taiwan
α	0.358 (5.443)	0.285 (4.193)		0.378 (5.228)	2.235 (10.843)
δ	0.7	0.5		1.1	1.1
			Level		
γ_i	-0.061 (4.281) {3}	-0.029 (2.330) {1}	-0.020 (2.158) {2}	$-0.334^{\dagger}$ (1.766) {1}	-0.061 (2.128) {2}
δ	0.5	0.5		1.5	1.0

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the (P)ARCH and (P)ARCH-L models.

†lag 2 of output growth (negative insignificant) also included.

The numbers in parentheses are z-Statistic.

The numbers in {} indicate the order of lags.

"power" coefficients $\delta = 1.5$. For South Korea and Taiwan, the estimated value of δ stay in a similar level. Various lags of growth (from 1 to 12) are considered with the best model chosen on the basis of the minimum value of the AIC. For all countries/regions, there is strong evidence that growth affects its uncertainty negatively. The estimated (absolute) "level" coefficient is in the range $0.334 < |\gamma_i| < 0.020$. This result is consistent with the theory outlined in Section 2, even though it does not prove that the channel outlined in Section 2 is in effect. More analysis of the mechanism through which growth affects its uncertainty is offered in the following section. Moreover, this finding is in broad agreement with the predictions of the analysis of Blackburn and Pelloni (2005) for a negative correlation between output growth and its variability. It should be kept in mind though that our evidence is in terms of causality rather than correlation.

Next, we report the estimation results of an AR-(P)ARCH-M model of growth with $g(h_{t)} = h_t$, $g(h_t) = \sqrt{h_t}$or $g(h_t) = ln(h_t)$ for all countries/regions. Table 3.43, 3.44 and 3.45 report the estimated parameters of interest for different form of variance in the mean equation, respectively. In Japan and Singapore the estimates for the "in-mean" parameter (k) are statistically insignificant no matter the form in which the time varying variance enters the specification of the mean (see the "Mean" panel of Table 3.43, 3.44 and 3.45). For South Korea, the effects are positive and significant at 5% $(g(h_t) = \sqrt{h_t})$

Table 3.43. (P)ARCH-ML Models, Conditional Variance in Mean

	Japan	South Korea		Singapore	Taiwan
		Mean			
k	0.026 (0.610)	0.034 (1.336)		-0.016 (0.703)	0.036 (2.616)
δ	0.7	0.7		1.1	1.2
		Mean-level			
k	0.010 (0.185)	0.021 (0.904)		-0.016 (0.808)	0.040 (3.645)
γ_i	-0.060 (4.182) $\{3\}$	-0.054 (2.210) $\{1\}$	-0.034 (2.023) $\{2\}$	$-0.146^{\dagger}$ (1.664) $\{1\}$	-0.166 (2.607) $\{2\}$
δ	0.5	0.7		1.2	1.3

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the various (P)ARCH-M and (P)ARCH -ML models. In all cases $g(h_t) = h_t$.
†lag 2 of output growth (negative insignificant) also included.
The numbers in parentheses are z-Statistic.
The numbers in {} indicate the order of lags.

and 1% ($g(h_t) = ln(h_t)$) levels. For Taiwan, the effect is positive and significant at the 1% level when conditional variance incorporated in the mean equation.

Finally, Table 3.43, 3.44 and 3.45 report the estimation results of an AR-(P)ARCH-ML model. That is, we estimate a system of equations that allows not only the current value of the conditional variance to affect average growth, but also allows lag of the latter to influence the former (see the "Mean-Level" panel of Table 3.43, 3.44 and 3.45). All "level" estimated coefficients are significant and robust to the inclusion or exclusion of "in-mean" effects. As with the (P)ARCH-L model, we again find support for the theory outlined in Section 2 in all countries/regions. The "level" parameter is in the range $0.166 < |\gamma_i| < 0.013$. Moreover, we find evidence of positive impact from growth uncertainty on growth itself for South Korea($g(h_t) = \sqrt{h_t}$) and Taiwan ($g(h_t) = h_t$ or $ln(h_t)$) at 5% significance level, which support Mirman (1971), Black (1987) and Blackburn (1999). Again, the "in-mean" parameter is insignificant in Japan and Singapore. In all countries/regions, the values of the "power" coefficients are below 1.3.

Table 3.44. (P)ARCH-ML Models, Conditional Standard Deviation in Mean

	Japan	South Korea	Singapore	Taiwan
		Mean		
k	0.15 (1.144)	0.257 (2.061)	−0.221 (0.873)	0.146 (1.352)
δ	0.7	0.6	1.1	1.1
		Mean-level		
k	0.094 (0.519)	0.198 (2.253)	0.327 (1.180)	0.117 (1.057)
γ_i	−0.075 (4.255) {3}	−0.019 (2.305) {1}; −0.013 (1.969) {2}	$-0.114^{\dagger}$ (1.683) {1}	−0.068 (1.701) {2}
δ	0.6	0.5	1.1	1.1

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the various (P)ARCH-M and (P)ARCH-ML models. In all cases $g(h_t) = ln(h_t)$.
†lag 2 of output growth (negative insignificant) also included. The numbers in parentheses are z-Statistic.
The numbers in {} indicate the order of lags.

Table 3.45. (P)ARCH-ML Models, Logarithm of the Conditional Variance in Mean

	Japan	South Korea	Singapore	Taiwan
		Mean		
k	0.083 (0.914)	0.574 (2.243)	−0.855 (1.022)	0.250 (1.254)
δ	0.7	0.5	1.1	1.0
		Mean-level		
k	0.131 (1.040)	0.355 (1.752)	−0.956 (1.162)	0.407 (5.107)
γ_i	−0.060 (4.204) {3}	−0.027 (2.116) {1}; −0.019 (2.003) {2}	$-0.116^{\dagger}$ (1.683) {1}	−0.134 (11.272) {2}
δ	0.5	0.5	1.1	1.2

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the various (P)ARCH-M and (P)ARCH-ML models. In all cases $g(h_t) = ln(h_t)$.
†lag 2 of output growth (negative insignificant) also included.
The numbers in parentheses are z-Statistic.
The numbers in {} indicate the order of lags.

3.6 Robustness Checks

To check the sensitivity of our results to the impact of inflation[66] on growth and output uncertainty, lags of inflation are included in both mean and variance equations, since no significant effect from inflation was detected in the variance, we reestimated these model with lags of inflation in mean only. Table 3.46 shows that inflation has significant negative impact on output growth in all countries/regions, which is broadly in line with existing empirical studies. Consisting with our previous results, higher growth lead to lower growth variation.

Table 3.46. (P)ARCH Models with Inflation

	Japan	South Korea	Singapore	Taiwan
		ARCH		
θ_π	-0.199 (2.108) {2}	-0.441 (10.439) {6}	-1.620 (2.163) {9}	-1.68 (2.584) {7}; -0.214 (3.261) {9}
α	0.369 (5.737)	0.351 (7.250)	0.386 (5.286)	0.663 (5.486)
δ	0.7	0.5	1.1	1.2
		Level		
θ_π	-0.184 (2.671) {2}	-0.417 (5.207) {6}	-1.713 (2.234) {9}	-0.176 (2.012) {7}; -0.212 (5.162) {9}
γ_i	-0.062 (4.196) {3}	-0.029 (2.392) {1}; -0.018 (1.787) {2}	$-0.118^{\dagger}$ (1.777) {1}	-0.046 (1.940) {2}
δ	0.5	0.5	1.1	0.9

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the (P)ARCH and (P)ARCH-L models. θ_π captures the effect of inflation on growth.
†lag 2 of output growth (negative insignificant) also included.
The numbers in parentheses are z-Statistic.
The numbers in {} indicate the order of lags.

Table 3.47, 3.48 and 3.49 report the estimation results of (P)ARCH-M and (P)ARCH-ML models of growth with inflation in mean, with $g(h_{t)} = h_t$, $g(h_t) = \sqrt{h_t}$and $g(h_t) = ln(h_t)$, respectively. for all countries/regions. We find evidence regarding the direction

66 Inflation is measured by growth of CPI.

Table 3.47. (P)ARCH-ML Models with Inflation, Conditional Variance in Mean

	Japan	South Korea		Singapore	Taiwan	
Mean						
k	0.052 (1.199)	0.043 (1.374)		−0.019 (1.058)	0.035 (3.033)	
θ_π	−0.202 (2.216) {2}	−0.400 (2.315) {6}		−1.659 (2.428) {9}	−0.219 (2.445) {7}	−0.195 (2.545) {9}
δ	0.7	0.6		0.9	1.6	
Mean-level						
k	0.037 (0.746)	0.044 (1.753)		−0.016 (0.840)	0.038 (3.672)	
θ_π	−0.032 (0.436) {2}	−0.486 (3.431) {6}		−1.692 (2.262) {9}	−0.207 (2.304) {7}	−0.190 (2.446) {9}
γ_i	−0.057 (3.865) {3}	−0.054 (2.210) {1}	−0.034 (2.023) {2}	$-0.084^\dagger$ (1.688) {1}	−0.226 (2.708) {2}	
δ	0.5	0.6		1.0	1.5	

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the various (P)ARCH-M and (P)ARCH-ML models. In all cases $g(h_t) = h_t$.
θ_π captures the effect of inflation on growth.
† lag 2 of output growth (negative insignificant) also included.
The numbers in parentheses are z-Statistic.
The numbers in {} indicate the order of lags.

of the impact of a change in uncertainty on growth. That is, in South Korea and Taiwan, volatility of growth has positive and significant impact on growth. In detail, we find that growth uncertainty positively affect growth in form of the logarithm of conditional variance(at 1% significance level) in South Korea, and in form of the conditional variance(at 1% significance level) and conditional standard deviation(at 10% significance level) in Taiwan, these results do not alter with impact of "level- effect". Again, "level- effect" and inflation impact are both negative.

Since we did not find any impact of inflation on growth uncertainty, in order to find out the channel that growth affects its uncertainty, we test the effect of the inflation uncertainty[67] on growth uncertainty. "GROWTH" panel of Table 3.50 shows that effects of

[67] We use conditional variance derivate from a (P)ARCH model of inflation as proxy of inflation uncertainty.

Table 3.48. (P)ARCH-ML Models with Inflation, Conditional Standard Deviation in Mean

	Japan	South Korea	Singapore	Taiwan
		Mean		
k	0.052 (0.417)	0.302 (3.232)	−0.366 (1.532)	0.224 (1.732)
θ_π	−0.198 (2.183) {2}	−0.545 (9.600) {6}	−1.647 (2.371) {9}	−0.190 (2.080) {7} −0.212 (3.091) {9}
δ	0.7	0.5	1.0	1.2
		Mean-level		
k	0.0002 (0.002)	0.178 (1.091)	−0.391 (1.528)	1.081 (1.929)
θ_π	−0.176 (2.413) {2}	−0.462 (2.783) {6}	−1.798 (2.434) {9}	−0.241 (1.636) {7} −0.189 (2.009) {9}
γ_i	−0.062 (4.329) {3}	−0.053 (2.290) {1} −0.032 (1.915) {2}	$-0.173^{\dagger}$ (2.026) {1}	−0.130 (2.166) {2}
δ	0.5	0.7	1.2	1.0

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the various (P)ARCH-M and (P)ARCH-ML models. In all cases $g(h_t) = \sqrt{h_t}$.

θ_π captures the effect of inflation on growth.

†lag 2 of output growth (negative insignificant) also included.

The numbers in parentheses are z-Statistic.

The numbers in {} indicate the order of lags.

Table 3.49. (P)ARCH-ML Models with Inflation, Logarithm of the Conditional Variance in Mean

	Japan	South Korea	Singapore	Taiwan
		Mean		
k	0.118 (1.054)	0.602 (3.475)	-0.689 (1.176)	0.210 (0.922)
θ_π	-0.185 (1.988) $\{2\}$	-0.467 (3.191) $\{6\}$	-1.698 (2.486) $\{9\}$	-0.181 (2.049) $\{7\}$ -0.207 (3.243) $\{9\}$
δ	0.7	0.6	0.9	1.0
		Mean-level		
k	0.100 (0.871)	0.432 (2.075)	-1.024 (1.414)	0.620 (5.861)
θ_π	-0.029 (0.387) $\{2\}$	-0.566 (3.369) $\{6\}$	-1.795 (2.461) $\{9\}$	-2.000 (2.049) $\{7\}$ -0.227 (2.887) $\{9\}$
γ_i	-0.058 $(4.037)g$ $\{3\}$	-0.047 (2.078) $\{1\}$ -0.027 (1.524) $\{2\}$	$-0.092^{\dagger}$ (1.814) $\{1\}$	-0.202 (5.757) $\{2\}$
δ	0.5	0.7	1.0	1.4

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the various (P)ARCH-M and (P)ARCH-ML models. In all cases $g(h_t) = ln(h_t)$.

θ_π captures the effect of inflation on growth.

† lag 2 of output growth (negative insignificant) also included.

The numbers in parentheses are z-Statistic.

The numbers in {} indicate the order of lags.

inflation uncertainty on growth uncertainty are negative and significant in Japan, South Korea, Taiwan and Singapore. Again, inflation has negative impact on growth in all countries/regions. Table 3.50 also presents a (P)ARCH model of inflation with lagged output growth variable been added in both mean and variance equations(see "INFLATION" panel of Table 3.50). In Japan, South Korea and Taiwan, it is very clear that growth has positive effect on both inflation and inflation uncertainty, however, no evidence of positive/negative effect of growth on inflation/inflation uncertainty was found in Singapore. In addition, the "level- effect" for inflation are positive in all countries/regions, which indicate that higher inflation uncertainty associate with higher inflation rate. Hence, combining the two panels in Table 3.50, we find in three out of four countries/regions that growth negatively impact on its uncertainty via the inflation uncertainty channel: higher growth leads to higher inflation uncertainty (directly and via "level- effect"), and higher inflation uncertainty induces lower growth uncertainty.

In order to capture the difference between low/negative growth periods (oil shock, demand shock, financial crisis, et al.) and normal growth period in the series, we use the methodology developed by Bai and Perron (2003) to examine whether there is any structural break in growth and when it occurred. Bai and Perron (2003) address the problem of testing for multiple structural changes under very general conditions on the data and the errors. In addition to testing for the existence of breaks, these statistics identify the number and location of multiple breaks. Table 3.51 reports the implicated break points and the corresponding economic events: DO stands for oil shock in 1970s, DH stands for the Japanese assets price bubble's burst in 1991, DF stands for the 1997 financial crisis in Asia, DE stands for the deep decline in demand for electronic industry in 2001, and DC is the recent credit crunch in 2008.

Table 3.52 reports the (P)ARCH models with structural break dummy variable incorporated in the mean and variance equations and with or without "level- effect". We

$\delta = 0.5, 0.5, 0.5$ and 0.6 for Japan, South Korea, Singapore and Taiwan respectively.

Table 3.50. (P)ARCH Models with Inflation Uncertainty

	Japan	South Korea	Singapore	Taiwan	
			GROWTH		
θ_π	-0.360	-0.340	-1.690	-0.194	-0.218
	(4.376)	(2.227)	(2.033)	(2.483)	(15.016)
	$\{2\}$	$\{6\}$	$\{9\}$	$\{7\}$	$\{9\}$
$\gamma_{h\pi}$	-0.291	-0.263	-6.485	-0.097	
	(3.301)	(1.818)	(1.928)	(7.874)	
	$\{2\}$	$\{8\}$	$\{4\}$	$\{8\}$	
δ	0.5	0.7	1.1	1.0	
			INFLATION		
θ_y	0.009	0.006	$-$	0.016	
	(2.339)	(1.734)		(1.884)	
	$\{5\}$	$\{6\}$		$\{6\}$	
γ_i	0.029	0.080	0.081	0.135	
	(2.315)	(3.227)	(2.133)	(3.093)	
	$\{1\}$	$\{1\}$	$\{1\}$	$\{1\}$	
γ_y	0.009	0.007	$-$	0.087	
	(5.620)	(2.564)		(3.056)	
	$\{5\}$	$\{5\}$		$\{5\}$	
δ	0.5	0.5	0.6	0.9	

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the (P)ARCH models.
$\theta_\pi(\theta_y)$ captures the effect of inflation (growth) on growth (inflation).
$\gamma_{h\pi}(\gamma_y)$ captures the effect of inflation uncertainty (growth) on growth uncertainty (inflation uncertainty).
The numbers in parentheses are z-Statistic.
The numbers in {} indicate the order of lags.

Table 3.51. Break Points and Dummy Variables

	DO	DH	DF	DE	DC
JAPAN	11/1973-03/1975	05/1991-03/1993		12/2000-12/2001	02/2008-09/2008-02/2009†
SOUTH KOREA			10/1997-07/1998		05/2008-01/2009
SINGAPORE	02/1974-12/1974			02/2001-12/2001	04/2008-01/2009
TAIWAN				11/2000-08/2001	03/2008-01/2009

DO: oil shock in1970s, DH: Japanese assets price bubble burst in 1991, DF: the1997 financial crisis in Asia, DE: the deep decline in demand for electronic industry in 2001, DC: the credit crunch in 2008.
The dummy variables equal to "1" in the corresponded period (start date not included) and "0" elsewhere.
†Japan has 2 dummy variables for credit crunch: DC1(01/2008-09/2008) and DC2(10/2008-02/2009)

find that dummies in all four countries/regions are statistically significant in the mean equation with negative sign and are irrelevant with "level- effect" incorporated or not, but all dummies are insignificant in the variance equation for South Korea, Singapore and Taiwan, we reestimated the models with dummies in mean only for these countries/regions. "level- effect" are negative in Japan, South Korea, Singapore(insignificant) and Taiwan. the IC chooses a (P)ARCH model with estimated power term parameter of $\delta = 0.7, 0.9, 0.6$ and 0.8 for Japan, South Korea, Singapore and Taiwan, respectively. The corresponding values for the (P)ARCH-L model are $0.6, 0.8, 0.9$ and 1.0.

Next, we report the estimation results of (P)ARCH-ML and (P)ARCH-M models of growth with structural break dummy variable incorporated in the mean and variance equation and $g(h_t) = h_t$, $\sqrt{h_t}$ and $ln(h_t)$ for all countries/regions. Table 3.53 3.54 and 3.55 report only the estimated parameters of interest and in all cases, the coefficients for the dummy variables are negative and significant. Regarding to "level- effect", the results are very similar to those obtained without using dummy variables, The growth of all countries/regions' have negative impact on its uncertainty, the strong evidence in all countries/regions is invariant to the form of the "in-mean" variable. Regarding the reverse causal effect, our evidence is country/region specific. As before, the estimates for the "in-mean" parameter (k) in South Korea and Taiwan are positive, interestingly, with the join of dummy variables, these effects are significant in 10% level with all form of growth uncertainty we tested and do not alter with inclusion of "level- effect". Very interestingly, the "in-mean" effects for Japan become significant at 10% level in all "MEAN" panel, however when lags of growth were added into the variance equation, this effect disappeared except the case of $g(h_t) = \sqrt{h_t}$. In all cases, "in-mean" effects are insignificant for Singapore. Again, all power term parameters of δ are below 1.5.

We also estimated models with both inflation and dummy variables. In all case, the impact from inflation becomes insignificant and dummy variables still have negative significant influence on growth. We do not present these results since they are very similar to the results with dummy variables only.

Table 3.52. (P)ARCH Models with Dummies

	Japan‡	South Korea	Singapore	Taiwan
		Normal		
θ_{do}	-1.736 (6.876)	–	–	-2.397 (10.668)
θ_{dh}	-1.324 (3.819)	–	–	–
θ_{df}	–	-5.203 (2.990)	–	–
θ_{de}	-1.498 (6.801)	–	-7.896 (3.736)	-1.295 (3.318)
θ_{dc}	DC1 DC2 -2.284 -7.622 (3.428) (6.432)	-3.520 (2.707)	-6.279 (5.302)	-7.481 (2.339)
α	0.223 (3.460)	0.418 (5.704)	0.313 (4.409)	0.561 (6.342)
δ	0.7	0.9	1.0	0.8
		Level		
θ_{do}	-1.653 (7.532)	–	–	-2.548 (4.191)
θ_{dh}	-1.219 (4.482)	–	–	–
θ_{df}	–	-5.309 (3.491)	–	–
θ_{de}	-1.514 (5.507)	–	-7.737 (3.715)	-1.109 (1.972)
θ_{dc}	DC1 DC2 -2.442 -7.104 (3.571) (5.023)	-4.476 (2.537)	-6.372 (5.471)	-7.339 (2.316)
γ_i	-0.065 (4.528) $\{3\}$	-0.076 -0.040 (2.294) (1.937) $\{1\}$ $\{2\}$	$-0.028^{\dagger}$ (0.795) (1)	-0.053 (1.738) $\{2\}$
δ	0.6	0.8	0.9	1.0

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the (P)ARCH and (P)ARCH-L models.
θ_{di},$(i = o, h, f, e, c)$ are coefficients of dummy variables.
† lag 2 of output growth (negative insignificant) also included.
‡ dummies in the variance equation also included.
The numbers in parentheses are z-Statistic.
The numbers in {} indicate the order of lags.

Table 3.53. (P)ARCH-ML Models with Dummies, Conditional Variance in Mean

	Japan‡		South Korea		Singapore	Taiwan
	Mean					
k	0.204 (1.985)		0.059 (2.206)		-0.002 (0.047)	0.107 (3.055)
θ_{do}	-1.618 (4.608)		–		–	-3.390 (6.806)
θ_{dh}	-1.361 (3.811)		–		–	–
θ_{df}	–		-2.898 (4.596)		–	–
θ_{de}	-1.255 (4.022)		–		-7.844 (3.735)	-1.497 (1.975)
	DC1	DC2				
θ_{dc}	-3.022 (3.003)	-8.650 (5.955)	-4.122 (2.491)		-6.422 (5.543)	-8.790 (4.917)
δ	1.0		1.0		0.9	1.3
	Mean-level					
k	0.248 (1.612)		0.048 (1.946)		-0.001 (0.049)	0.140 (2.798)
θ_{do}	-1.382 (4.313)		–		–	-3.672 (7.000)
θ_{dh}	-1.184 (3.229)		–		–	–
θ_{df}	–		-4.737 (2.740)		–	–
θ_{de}	-1.100 (3.038)		–		-7.780 (3.697)	-4.581 (2.272)
	DC1	DC2				
θ_{dc}	-2.939 (3.004)	-9.174 (7.196)	-4.273 (2.435)		-6.255 (5.073)	-10.474 (8.842)
γ_i	-0.105 (3.943) $\{3\}$		-0.078 (1.735) $\{1\}$	-0.042 (1.731) $\{2\}$	$-0.038^{\dagger}$ (0.818) $\{1\}$	-0.087 (1.542) $\{2\}$
δ	0.5		0.9		1.0	1.3

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the various (P)ARCH-M and (P)ARCH-ML models. In all cases $g(h_t) = h_t$.
$\theta_{di}, (i = o, h, f, e, c)$ are coefficients of dummy variables.
† lag 2 of output growth (negative insignificant) also included.
‡ dummies in the variance equation also included.
The numbers in parentheses are z-Statistic.
The numbers in {} indicate the order of lags.

3.7 Conclusions

We use monthly data for four Asian countries/regions to test the relationship between output growth and its variability. Our empirical approach employs a PARCH model that allows lagged growth to appear in the conditional variance equation. Various alternative specifications are estimated in order to confirm the robustness of our results. These specifications allow for different forms of growth uncertainty as regressor in the mean, also allow the impact of inflation on growth and growth uncertainty. In addition, our models are also estimated with dummy variables in order to establish robustness to major events such as oil shocks and the credit crunch. We derive two main conclusions. First, we find evidence in supporting of Mirman(1971), Black (1987) and Blackburn (1999) that output volatility has positive effect on economic performance in three of the four countries/regions, however, significance of this positive impact is subjects to the form of growth uncertainty, we find that, growth uncertainty in form of the logarithm of conditional variance in South Korea and in form of conditional variance in Taiwan are most robust. For Japan, we find significant positive impact from growth uncertainty on growth when dummy variables were taken in to account but only when the growth uncertainty take the form of conditional standard deviation can survive in the (P)ARCH-ML models. Second, we find strong evidence that growth negatively affects its uncertainty via the channel of inflation uncertainty. Overall, the relationship between growth and its variability which, in most of the cases considered, is bi-directional (with the causal effect of the former on the latter being negative). These results support the recent emphasis on the treatment of the variability of the business cycle in tandem with the theory of economic growth. Furthermore, our evidence for bi-directional causality between growth and its variability with mixed signs concurs with the predictions of theoretical models by Blackburn and Galinder (2003) and Blackburn and Pelloni (2004) that either type of correlation (positive or negative) between the two variables depend on the type of learning and the type of shocks hitting the economy.

Table 3.54. (P)ARCH-ML Models with Dummies, Conditional Standard Deviation in Mean

	Japan[‡]	South Korea	Singapore	Taiwan
	Mean			
k	1.374 (1.859)	0.443 (2.411)	−0.021 (0.062)	0.669 (3.145)
θ_{do}	−2.173 (3.594)	–	–	−3.265 (6.811)
θ_{dh}	−1.608 (3.354)	–	–	–
θ_{df}	–	−5.009 (2.497)	–	–
θ_{de}	−1.251 (2.659)	–	−7.751 (3.729)	−1.258 (1.505)
θ_{dc}	DC1: −3.236 (2.465); DC2: −11.785 (4.776)	−4.266 (2.438)	−6.612 (5.765)	−8.762 (2.387)
δ	0.7	0.9	0.8	1.2
	Mean-level			
k	0.765 (1.701)	0.342 (1.852)	−0.033 (0.105)	2.017 (4.156)
θ_{do}	−1.333 (3.542)	–	–	−4.473 (10.612)
θ_{dh}	−1.206 (3.279)	–	–	–
θ_{df}	–	−3.292 (5.288)	–	–
θ_{de}	−1.007 (2.503)	–	−7.661 (3.725)	−4.974 (1.994)
θ_{dc}	DC1: −2.843 (3.085); DC2: −8.294 (7.600)	−4.923 (2.602)	−6.585 (5.682)	−10.350 (3.872)
γ_i	−0.073 (4.419) {3}	−0.048 (1.914) {1}; −0.029 (1.728) {2}	$-0.019^{\dagger}$ (0.738) {1}	−0.182 (8.302) {2}
δ	0.7	0.7	0.8	1.4

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the various (P)ARCH-M and (P)PARCH-ML models. In all cases $g(h_t) = \sqrt{h_t}$. $\theta_{di}, (i = o, h, f, e, c)$ are coefficients of dummy variables.

†lag 2 of output growth (negative insignificant) also included.

‡dummies in the variance equation also included.

The numbers in parentheses are z-Statistic.

The numbers in {} indicate the order of lags.

Table 3.55. (P)ARCH-ML Models with Dummies, Logarithm of the Conditional Variance in Mean

	Japan‡	South Korea	Singapore	Taiwan
	Mean			
k	0.386 (1.838)	0.612 (2.503)	-0.072 (0.088)	1.125 (3.038)
θ_{do}	-1.605 (4.186)	–	–	-3.225 (6.418)
θ_{dh}	-1.421 (4.023)	–	–	–
θ_{df}	–	-2.557 (2.995)	–	–
θ_{de}	-1.178 (3.599)	–	-7.627 (3.683)	-1.168 (1.438)
θ_{dc}	DC1: -2.707 (2.873); DC2: -7.951 (6.213)	-4.520 (2.773)	-6.713 (6.008)	-8.589 (2.358)
δ	1.0	0.8	0.8	1.1
	Mean-level			
k	0.314 (1.248)	0.486 (2.330)	-0.059 (0.069)	2.753 (5.778)
θ_{do}	-1.496 (4.004)	–	–	-4.875 (10.382)
θ_{dh}	-1.284 (3.463)	–	–	–
θ_{df}	–	-3.173 (6.027)	–	–
θ_{de}	-1.164 (2.788)	–	-7.754 (3.727)	-4.271 (11.471)
θ_{dc}	DC1: -2.590 (3.083); DC2: -7.648 (6.133)	-4.935 (2.626)	-6.428 (5.391)	-10.222 (2.948)
γ_i	-0.097 (4.388) $\{3\}$	-0.035 (1.934) $\{1\}$; -0.021 (1.491) $\{2\}$	$-0.026^{\dagger}$ (0.747) $\{1\}$	-0.139 (10.796) $\{2\}$
δ	0.9	0.6	0.9	1.3

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the various (P)ARCH-M and (P)ARCH-ML models. In all cases $g(h_t) = ln(h_t)$. θ_{di},$(i = o, h, f, e, c)$ are coefficients of dummy variables.

† lag 2 of output growth (negative insignificant) also included.

‡ dummies in the variance equation also included.

The numbers in parentheses are z-Statistic.

The numbers in {} indicate the order of lags.

3.A Appendix

Table 3.56. Power Term Parameters of (P)ARCH-ML Models

	Japan	South Korea	Singapore	Taiwan
Simple	0.7	0.5	1.1	1.1
Level	0.5	0.5	1.5	1.0
With Inflation				
Simple	0.7	0.5	1.1	1.2
Level	0.5	0.5	1.1	0.9
With Structural Break				
Simple	0.7	0.9	1.0	0.8
Level	0.6	0.8	0.9	1.0

Notes: For each of the four Asian countries/regions, this table reports the power term Parameters of the various (P)ARCH and (P)ARCH-L models.

Table 3.57. Power Term Parameters of (P)ARCH-ML Models with Conditional Variance in Mean

	Japan	South Korea	Singapore	Taiwan
Mean	0.7	0.7	1.1	1.2
Mean-level	0.5	0.7	1.2	1.3
With Inflation				
Mean	0.7	0.6	0.9	1.6
Mean-level	0.5	0.6	1.0	1.5
With Structural Break				
Mean	1.0	1.0	0.9	1.3
Mean-level	0.5	0.9	1.0	1.3

Notes: For each of the four Asian countries/regions, this table reports the power term Parameters of the various (P)ARCH-M and (P)ARCH-ML models. In all cases $g(h_t) = h_t$.

Table 3.58. Power Term Parameters of (P)ARCH-ML Models with Conditional Standard Deviation in Mean

	Japan	South Korea	Singapore	Taiwan
Mean	0.7	0.6	1.1	1.1
Mean-level	0.6	0.5	1.1	1.1
With Inflation				
Mean	0.7	0.5	1.0	1.2
Mean-level	0.5	0.7	1.2	1.0
With Structural Break				
Mean	0.7	0.9	0.8	1.2
Mean-level	0.7	0.7	0.8	1.4

Notes: For each of the four Asian countries/regions, this table reports the power term Parameters of the various (P)ARCH-M and (P)ARCH-ML models. In all cases $g(h_t) = \sqrt{h_t}$.

Table 3.59. Power Term Parameters of (P)ARCH-ML Models with Logarithm of the Conditional Variance in Mean

	Japan	South Korea	Singapore	Taiwan
Mean	0.7	0.5	1.1	1.0
Mean-level	0.5	0.5	1.1	1.2
With Inflation				
Mean	0.7	0.6	0.9	1.0
Mean-level	0.5	0.7	1.0	1.4
With Structural Break				
Mean	1.0	0.8	0.8	1.1
Mean-level	0.9	0.6	0.9	1.3

Notes: For each of the four Asian countries/regions, this table reports the power term Parameters of the various (P)ARCH-M and (P)ARCH-ML models. In all cases $g(h_t) = ln(h_t)$.

Chapter 4
Is the Relationship Between Inflation and its Uncertainty Linear? Evidence from Asia

4.1 Introduction

[68]The issue of the welfare costs of inflation has been one of the most researched topics in macroeconomics both on the theoretical and empirical fronts. Friedman (1977) argues that a rise in inflation leads to more nominal uncertainty. The opposite direction of causation has also been analyzed in the theoretical literature. Cukierman and Meltzer (1986) argue that central banks tend to create inflation surprises in the presence of more nominal uncertainty. Clarida et al. (1999) emphasize the fact that since the late 1980s a stream of empirical work has presented evidence that monetary policy may have important effects on real activity. Consequently, there has been a great resurgence of interest in the issue of how to conduct monetary policy. If an increase in the rate of inflation causes an increase in its uncertainty, one can conclude that greater uncertainty - which many have found to be negatively correlated to economic activity - is part of the costs of inflation. Thus, if we attempt to provide a satisfactory answer to the questions “What actions should central bankers take?” and “What is the optimal strategy for monetary authorities to follow?” we must first develop a clear view about the temporal ordering of inflation and nominal uncertainty.

Many studies examining the link between inflation and uncertainty use GARCH estimation methods, but they differ in the choice of sample periods, frequency data sets and empirical specifications. For example, Baillie et al. (1996) employ an ARFIMA-

[68] This chapter is a sister paper of Karanasos and Schurer (2008), “Is the Relationship Between Inflation and its Uncertainty Linear?” Introduction and literature sections are generally based on Karanasos and Schurer (2008).

GARCH-in-mean model, Grier and Perry (1998) and Fountas and Karanasos (2007) estimate univariate component GARCH specifications, Conrad and Karanasos (2005a, b) utilize the ARFIMA-FIGARCH model, and Fountas et al. (2006) use a bivariate constant correlation GARCH formulation.

Despite using different GARCH specifications, all these studies focus exclusively on the standard Bollerslev type of model which assumes the conditional variance is a linear function of lagged squared errors. There seems to be, however, no economic reason why one should make such a strong assumption. The common use of a squared term in this role is most likely to be a reflection of the normality assumption traditionally invoked working with inflation data. However, if we accept that inflation data are very likely to have a non-normal error distribution, then the superiority of a squared term is lost and other power transformations may be more appropriate. Indeed, for non-normal data, by squaring the inflation rates one effectively imposes a structure on the data which may potentially furnish sub-optimal modeling and forecasting performance relative to other power terms. If π_t represents inflation in period t, Considers the temporal properties of the functions of $|\pi_t|^d$ for positive values of d. We find, as an empirical fact, that the autocorrelation function of $|\pi_t|^d$ is a concave function of d and reaches its maximum when d is smaller than one. This result serves as an argument against a Bollerslev type model.

In this chapter, we illustrate these concerns empirically for Japan, South Korea, Singapore and Taiwan using a parametric power ARCH model (PARCH). The PARCH model can be viewed as a standard GARCH model for observations that have been changed by a sign-preserving power transformation implied by a (modified) PARCH parameterization. The PARCH model increases the flexibility of the conditional variance specification by allowing the data to determine the power of inflation for which the predictable structure in the volatility pattern is the strongest. This feature in the volatility processes of inflation has major implications for the inflation-uncertainty hypothesis. To test for the relationship between the two variables we use the simultaneous-estimation approach. Under this approach, we estimate a PARCH-in-mean model with the conditional variance equation incorporating lags of the inflation series (the “level” effect), thus allowing simultaneous estimation and testing of the bidirectional causality between the inflation series and the associated uncertainty. Moreover, He and Terävirta (1999) emphasize that if the standard Bollerslev type of model is augmented by the ‘heteroscedasticity’ parameter (the “power” term), the estimates of the ARCH and GARCH coefficients almost certainly change. More importantly, we find that the inflation-uncertainty relationship is sensitive to changes in the values of the “heteroscedasticity” parameter. Put differently, the estimated values of the “in-mean” and the “level” effects are fragile to changes in the “power” term.

The chapter is organized as follows: In section 2 we consider in more detail the hypotheses about the causality between inflation and its uncertainty. In Section 3, we describe the time series model for inflation and explain its merits. We report the empirical results in Section 4 and in Section 5 we evaluate the robustness of our findings. Section 6 outlines our conclusions and discusses our results.

4.2 The Link between Inflation and Its Uncertainty

4.2.1 Theory

The effect of inflation on its uncertainty is theoretically ambiguous. The Friedman (1977) hypothesis stresses the harmful effects of nominal uncertainty on employment and production. On this basis several researchers contend that a high rate of inflation produces greater uncertainty about the future direction of government policy and, thus, about future rates of inflation. Ball (1992) formalizes this idea in the context of a repeated game between the monetary authority and the public. This extension of a Barro-Gordon model introduces exogenous shocks and two Central Bank (CB) policy-makers, one Conservative and one Liberal, who have different preferences over how to react in times of high inflation. During these times the public is confused because it does not know which policy maker is in charge. This incomplete information, in return, increases the public's uncertainty about future inflation. In accordance with the Friedman hypothesis we test for a positive effect.

In contrast, Ungar and Zilberfarb (1993) propose a mechanism that may weaken, offset, or even reverse the direction of the traditional view concerning the inflation-uncertainty relationship. They argue that, as inflation rises, economic agents invest more resources in forecasting it, thus reducing nominal uncertainty. However, this effect might only be present in periods of extreme inflation, which means that it comes into action only if the inflation rate surpasses a crucial threshold.

On the other hand, Cukierman and Meltzer (1986) predict that an increase in uncertainty will raise inflation due to the behavior of the CB in an uncertain environment. Their model is embedded in a Barro-Gordon setting in which the CB is not tied to a commitment rule on money supply growth. Therefore, the CB can pursue both objectives of "keeping inflation low" and "stimulating the economy by surprise inflation". Since the objective function of the CB and the money supply process are modelled as random variables, the public has difficulties inferring what caused higher inflation. It could be either that the CB finds it more important to stimulate the economy or that a random money supply shock occurred. Due to this information asymmetry the CB has an incentive to create inflation surprises in the presence of higher nominal uncertainty. In accordance with the Cukierman and Meltzer hypothesis, we test for a positive effect.

Finally, Holland (1995) predicts the opposite effect of uncertainty on inflation. He assumes the CB to be motivated by a desire for stability. If the CB analysts observe increasing nominal uncertainty due to an increasing inflation rate, the CB will restrict money supply growth. This measure is justified by reducing the potential of severe negative welfare effects of increasing inflation. In accordance with the Holland hypothesis, we test for a negative effect.

4.2.2 Empirical Evidence

The relationship between the two variables has been analyzed extensively in the empirical literature. Recent time series studies have focused particularly on the GARCH conditional variance of inflation as a statistical measure of nominal uncertainty (see, for example, Grier and Perry, 2000). To test for the relationship between uncertainty and indicators of macroeconomic performance such as inflation one can use either the two-step or the simultaneous-estimation approach.

Under the former approach, estimates of the conditional variance are obtained from the estimation of a standard GARCH model and these estimates are used in running Granger-causality tests to examine the causality between the two variables. Under the latter approach the model is estimated with the conditional variance (lagged inflation) included as a right-hand side variable in the mean (variance) equation.

Applying the two-step methodology, Grier and Perry (1998) in the G7 countries, Conrad and Karanasos (2005b) in Japan, UK and US, Thornton (2007) in some emerging market economies and Jiranyakul and Opiela (2010) in ASEAN-5 countries, find that inflation significantly raises its uncertainty. They also find evidence in favour of the Cukierman-Meltzer hypothesis for some countries and in favour of the Holland hypothesis for other countries. Their results regarding the impact of uncertainty on inflation were generally consistent with the rankings of CB independence (CBI).

Some studies use GARCH models that include a function of the lagged inflation rate in the conditional variance equation. In particular, Brunner and Hess (1993) allow for asymmetric effects of inflation shocks on nominal uncertainty and find a weak link between the two variables in the US. Two studies use GARCH type models with a joint feedback between the conditional mean and variance of inflation. Baillie et al. (1996), for three high inflation countries and the UK, and Karanasos et al. (2004) for the US, find strong evidence in favour of a positive bidirectional relationship in accordance with the predictions of economic theory. Karanasos and Schurer (2008) examine three European countries by using PARCH models. The paper showed evidence for Friedman hypothesis, but results for the effect of uncertainty on inflation are country specific.

4.3 PARCH Model

We follow the methodology of Karanasos and Schurer (2008) by adopting PARCH estimation for our inflation data. Since its introduction by Ding et al. (1993), the PARCH model has been frequently applied. For example, Hentschel (1995) defined a parametric family of asymmetric GARCH formulations that nests the EGARCH and PARCH models. He and Teräsvirta (1999) considered a family of first-order asymmetric GARCH processes which includes the asymmetric PARCH (A-PARCH) as a special case. Brooks et al. (2000) analyzed the applicability of the PARCH models to national stock market returns for ten countries[69]. Laurent (2004) derives analytical expressions for the score of the A-PARCH model. The use of the PARCH model is now widespread in the literature (see, for example, Mittnik and Paolella, 2000, Giot and Laurent, 2003, Karanasos and Schurer, 2005, Karanasos and Kim, 2006, and Conrad et al., 2007).

Let π_t follow an autoregressive (AR) process augmented by a "risk premium" defined in terms of volatility

$$\Phi(L)\pi_t = \phi_0 + kg(h_t) + \varepsilon_t \tag{4.22}$$

[69] It is worth noting that Fornari and Mele (1997) show the usefulness of the PARCH scheme in approximating models developed in continuous time as systems of stochastic differential equations. This feature of GARCH schemes has usually been overshadowed by their well-known role as simple econometric tools providing reliable estimates of unobserved conditional variances (Fornari and Mele, 2001).

with

$$\varepsilon_t = e_t h_t^{\frac{1}{2}}$$

where by assumption the finite order polynomial $\Phi(L) = \sum_{i=1}^{p} \phi L^i$ has zeros outside the unit circle. In addition, $\{e_t\}$ are independent and identically distributed (i.i.d) random variables with $E\left(e_t\right) = E\left(e_t^2 - 1\right) = 0$. The conditional variance of inflation $\{\pi_t\}$, is positive with probability one and is a measurable function of the sigma-algebra $\sum_{t-1}$, which is generated by $\{\pi_{t-1}, \pi_{t-2}, ...\}$.

Furthermore, we need to choose the form in which the time-varying variance enters the specification of the mean to determine the "risk premium". This is a matter of empirical evidence. In the empirical results that follow we employ three specifications for the functional form of the 'risk premium' ($g\left(h_t\right) = h_t$, $g\left(h_t\right) = \sqrt{h_t}$, or $g\left(h_t\right) = ln\left(h_t\right)$).

Moreover, h_t is specified as an A-PARCH(1, 1) process with lagged inflation included in the variance equation

$$h_t^{\frac{\delta}{2}} = \omega + \alpha f\left(e_{t-1}\right) + \beta h_{t-1}^{\frac{\delta}{2}} + \gamma_l \pi_{t-l} \tag{4.23}$$

with

$$f\left(e_{t-1}\right) = \left[\left|e_{t-1}\right| - \zeta e_{t-1}\right]^{\delta}$$

where δ with $\delta > 0$ is the 'heteroscedasticity' parameter, α and β are the ARCH and GARCH coefficients respectively, ζ with $|\zeta| < 1$ is the "leverage" term and γ_l is the "level" term for the lth lag of inflation. The model imposes a Box-Cox power transformation of the conditional standard deviation process and the asymmetric absolute residuals. The expected value of $f\left(e_{t-1}\right)$ is given by:

$$E\left[f\left(e_{t-1}\right)\right]=\begin{Bmatrix} \frac{1}{\sqrt{\pi}}\left[\left(1-\zeta\right)^{\delta}+\left(1+\zeta\right)^{\delta}\right]2^{\left(\frac{\delta}{2}-1\right)}\Gamma\left(\frac{\delta+1}{2}\right), \text{if } e_{t-1}\overset{(i.d)}{\sim}N(0,1) \\ \frac{(r-2)^{\frac{\delta}{2}}\Gamma\left(\frac{r-\delta}{2}\right)\Gamma\left(\frac{\delta+1}{2}\right)}{\Gamma\left(\frac{r}{2}\right)2\sqrt{\pi}}\left[\left(1-\zeta\right)^{\delta}+\left(1+\zeta\right)^{\delta}\right], \text{if } e_{t-1}\overset{(i.d)}{\sim}t_r(0,1) \end{Bmatrix}$$

where N and t denote the Normal and student-t distributions respectively, r are the degrees of freedom of the student-t distribution and $\Gamma(.)$ is the Gamma function. The δth moment of the conditional variance is a function of the above expression (see Karanasos and Kim, 2006).

Within the A-PARCH model, by specifying permissible values for δ, α, β, ζ and γ_l in Equation (4.23), it is possible to nest a number of the more standard ARCH and GARCH specifications (see Ding et al., 1993, Hentschel, 1995, and Brooks et al., 2000). For example, in Equation (4.23) let $\delta = 2$ and $\zeta = \gamma_l = 0$ to get the GARCH model. In order to distinguish the general model in Equations (4.22) and (4.23) from a version in which $k = \gamma_l = \zeta = \beta = 0$, we will hereafter refer to the former as A-PARCH-in-mean-level (A-PARCH-ML) and the latter as PARCH.

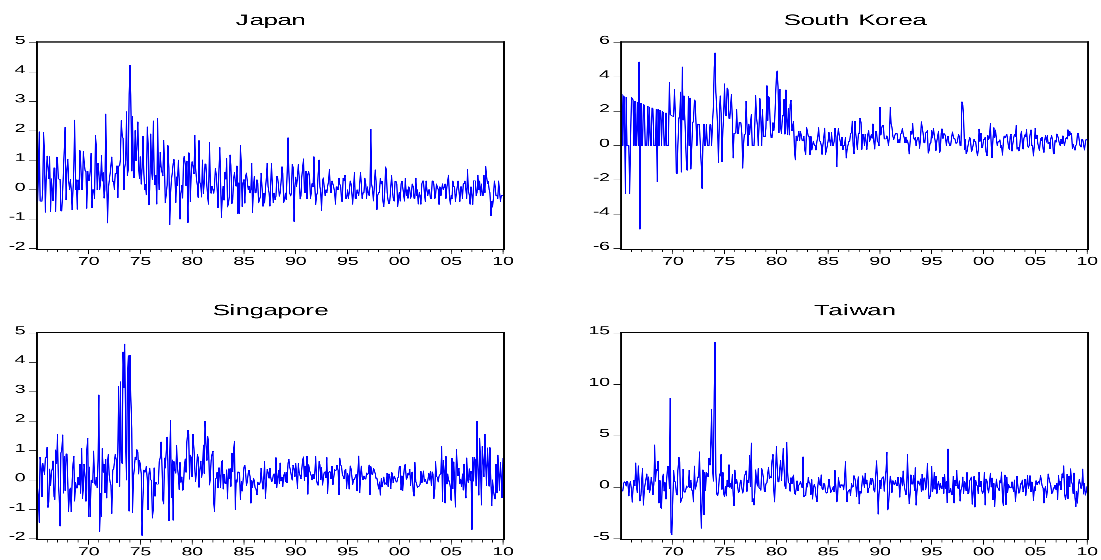
4.4 Empirical Analysis

4.4.1 Power-transformed Inflation

We use monthly data on the Consumer Price Index (CPI) as proxies for the price level[70]. The data range from 1965:02 to 2010:01 and covers four Asian countries/regions, namely, Japan, South Korea, Singapore and Taiwan.

[70] Most studies use CPI based inflation measures (i.e., Conrad and Karanasos, 2005a, b), therefore, we construct our measures also from the CPI. Alternatively, one can use either the Producer Price Index (PPI) or the GNP deflator. Brunner and Hess (1993) use all three measures but they discuss only the results using CPI inflation. Grier and Perry (2000) and Fountas and Karanasos (2007) use both (CPI and PPI) indices and find that the results are virtually identical.

Fig. 4.18. Inflation over Time



Inflation is measured by the difference between two months of the logarithm of CPI, i.e. $\pi = 100ln\left(CPI_t/CPI_{t-1}\right)$, which leaves us with 540 usable observations. The inflation rates of the four countries/regions are plotted in Figure 4.18. Japan, Singapore and Taiwan had a period of relatively high inflation during 1970s, this was mainly due to the 1970s' "oil shock". In South Korea, not only the "oil shock" but also the South Korea's "export-oriented growth strategy" during 1960s to 1970s lead to high inflation. this high inflation did not decrease until South Korean authorities started to hold down and stabilise money growth in late 1970s ((Dueker and Kim, 1999).

Figure 4.19 shows inflation by season for these four countries/regions, all of them revealed clear seasonal pattern. Census X12 seasonal adjustment is applied in order to remove the seasonal circle. Figure 4.20 and Figure 4.21 present the seasonal adjusted inflation. All tests and estimations in the rest of this chapter is based on seasonal adjusted data.

Fig. 4.19. Inflation by Season

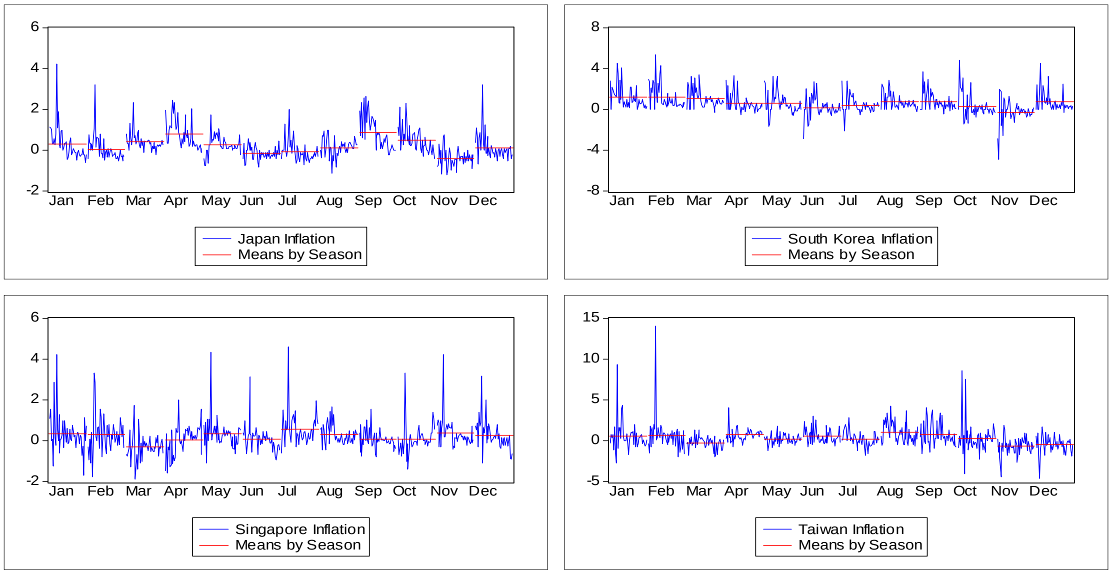


Fig. 4.20. Inflation over Time, Seasonal Adjusted

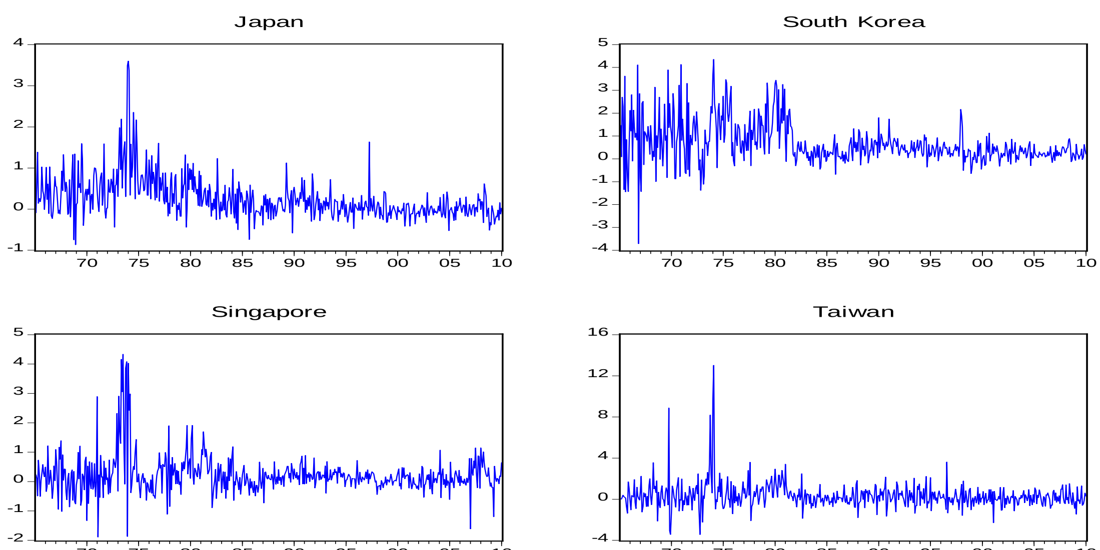
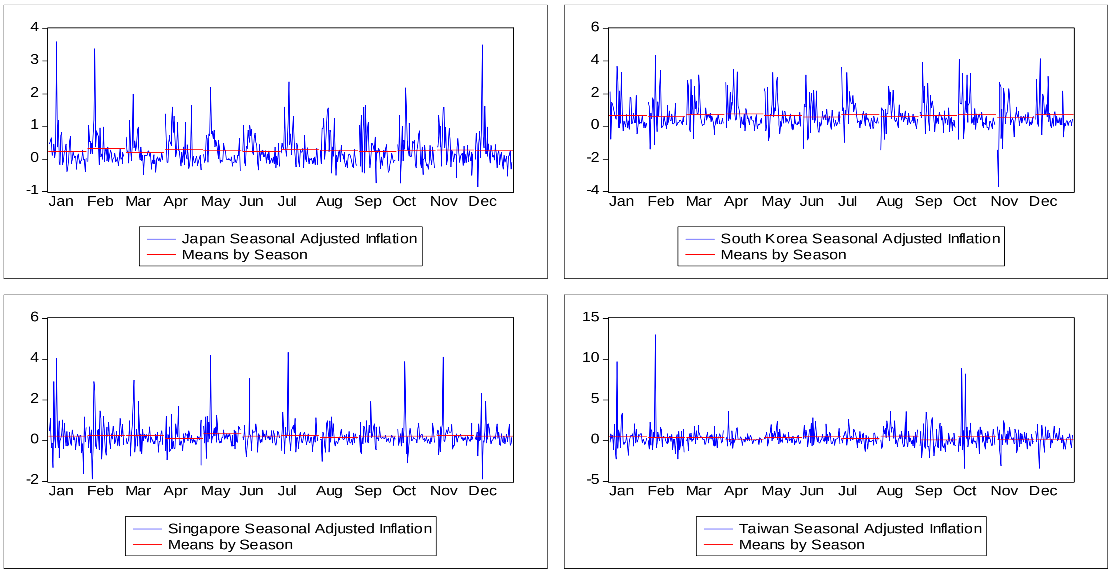


Fig. 4.21. Inflation by Season, Seasonal Adjusted

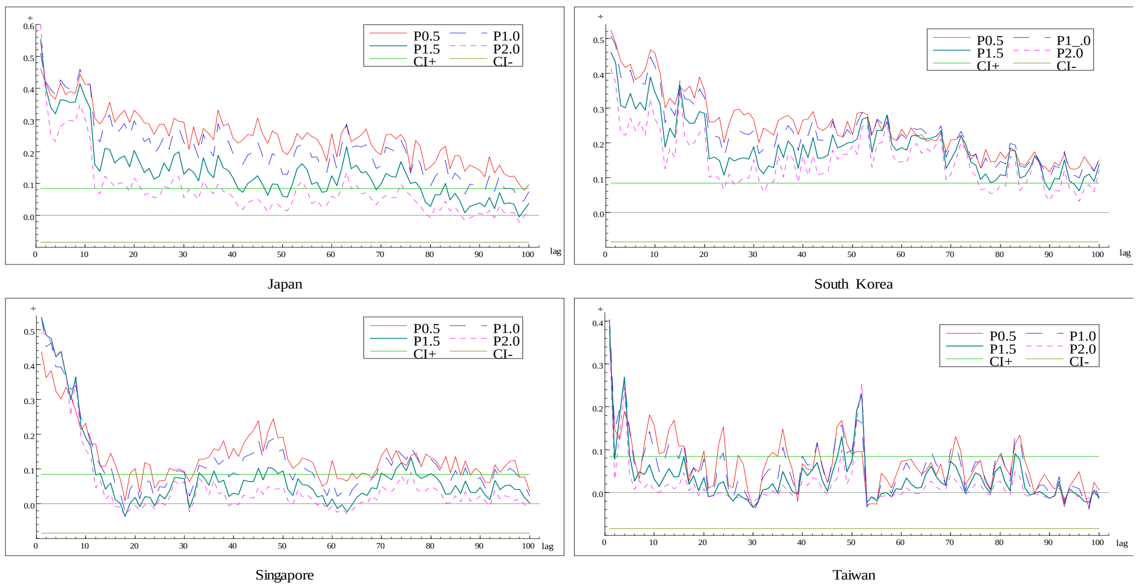


The results of the Phillips-Perron unit root tests (not reported) imply that we can treat the three rates as stationary processes. The summary statistics (not reported) indicate that the distribution of the three series is skewed to the right and has fat tails. The large values of the Jarque-Bera statistic imply a deviation from normality.

Next, we examine the sample autocorrelations of the power transformed absolute inflation $|\pi_t|^d$ for various positive values of d. Figure 4.22 shows the autocorrelogram of $|\pi_t|^d$ from lag 1 to 100 for $d = 0.5, 1.0, 1.5$ and 2.0. The horizontal lines show the $\pm 1.96/\sqrt{T}$ confidence interval (CI) for the estimated sample autocorrelations if the process π_t is i.i.d. In our case, $T = 540$, so CI $= \pm 0.0872$.

The sample autocorrelations for $\sqrt{|\pi_t|}$ are greater than the sample autocorrelations of $|\pi_t|^d$ for $d = 1.0, 1.5$ and 2.0, for most lags. In other words, the most interesting finding from the autocorrelogram is that $|\pi_t|^d$ has the strongest and slowest decaying autocorrelation when $d = 0.5$. Furthermore, the power transformations of absolute inflation when d is less than or equal to one have significant positive autocorrelations at least up to lag

Fig. 4.22. Autocorrelation of $|\pi_t|^d$



98 and 100 for Japan and South Korea. This power term is also produced most lags with significant positive autocorrelation for Singapore and Taiwan.

Figure 4.23 shows the autocorrelogram for power transformations of absolute residuals $|\varepsilon_t|^d$ from AR models (see Table 4.60). We plot the sample autocorrelations from lag 1 to 100 for $d = 0.5, 1.0, 1.5$ and 2.0. In general, the most interesting finding from the autocorrelogram is that, at most lags, $|\varepsilon_t|^d$ has the lowest autocorrelation when $d = 2.0$.

To illustrate this more clearly, we calculate the sample autocorrelations of the absolute value of inflation $\rho_\tau(d)$ as a function of d for lags $\tau = 12, 24$ and 36 and taking $d = 0.1, 0.2, 0.3$, to 4.5. Figure 4.24 provides the plots of the calculated $\rho_\tau(d)$. For example, for lag 12, there is a unique point d^* equal to $0.5, 0.5, 0.4$ and 0.5 for Japan, South Korea, Singapore and Taiwan respectively, such that $\rho_{12}(d)$ reaches its maximum at this point: for $\rho_{12}(d^*) > \rho_{12}(d)$ for $d^* \neq d$.

Fig. 4.23. Autocorrelation of $|\varepsilon_t|^d$

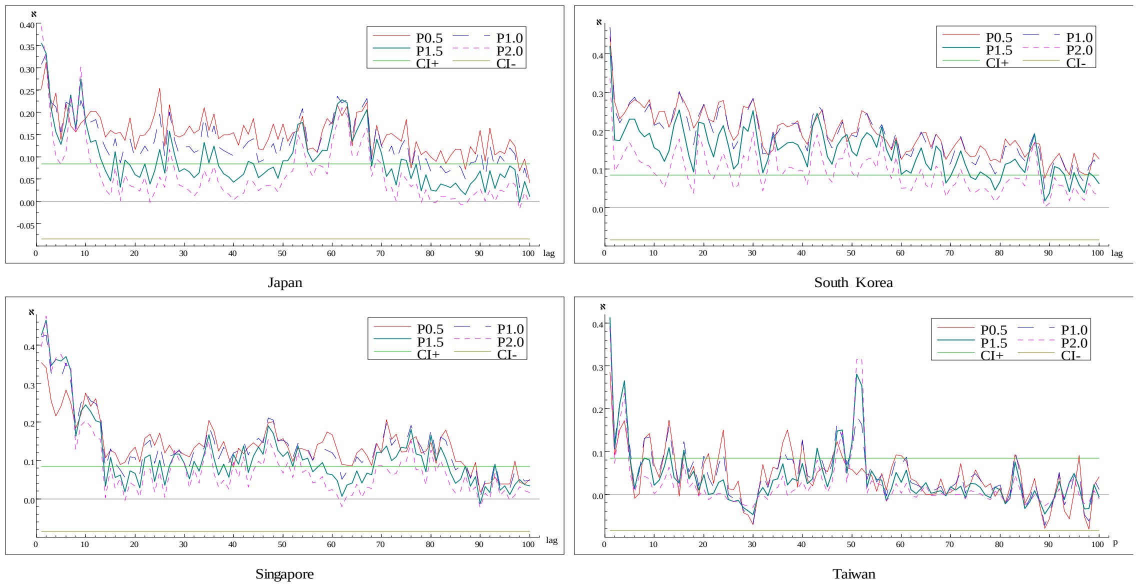


Fig. 4.24. Autocorrelation of $|\pi_t|^d$ at Lag 12, 24, 36

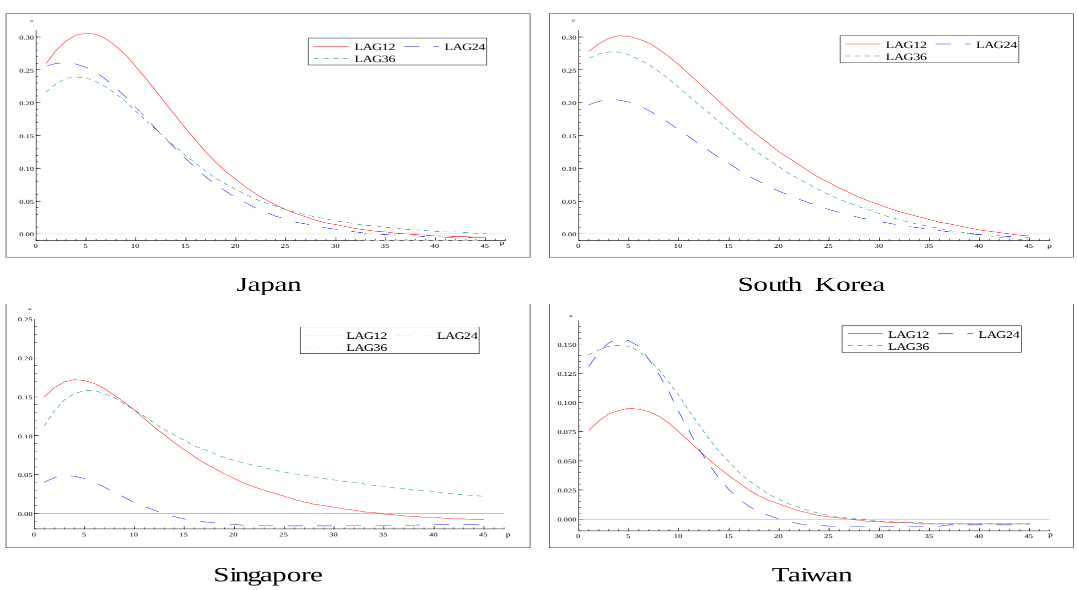
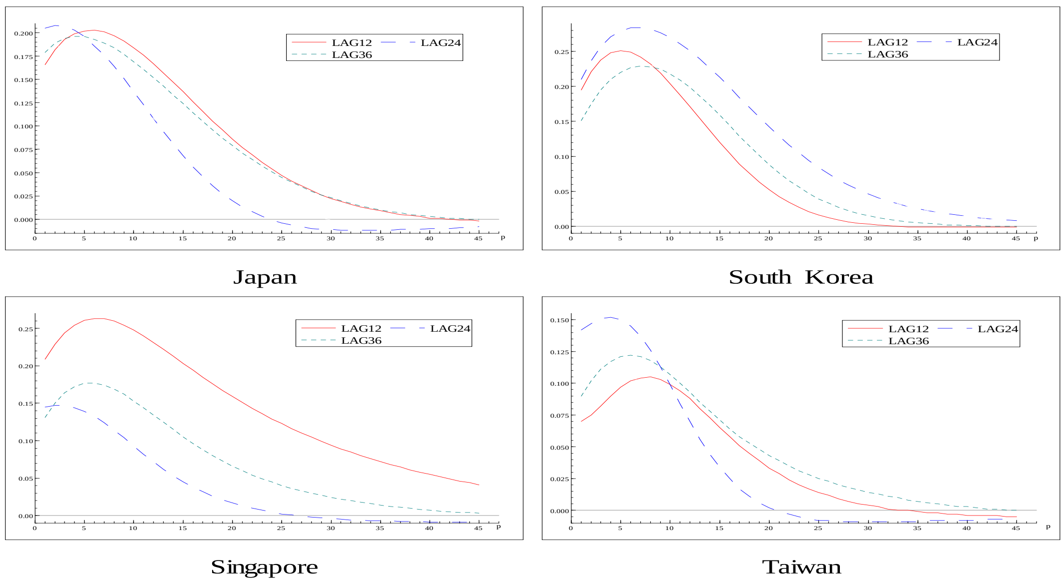


Fig. 4.25. Autocorrelation of $|\varepsilon_t|^d$ at Lag 12, 24, 36



Since for the choice of the econometric model it is important whether the strength of autocorrelation persists in the residuals of the model, we analogously present in Figure 4.25 the plots of calculated $\rho_\tau(d)$ for $|\varepsilon_t|^d$. For example, $\rho_{12}(d)$, reaches its maximum at $0.6, 0.6, 0.7$ and 0.7 for Japan, South Korea, Singapore and Taiwan respectively. These figures confirm the claim that in our data the autocorrelation structure of inflation is the strongest for values of d smaller than one.

4.4.2 Estimated Models of Inflation

We proceed with the estimation of the AR-PARCH(1,1) model in Equations (4.22) and (4.23) in order to take into account the serial correlation observed in the levels and power transformations of our time series data. These were obtained by quasi-maximum likelihood estimation (QMLE) as implemented in EVIEWS. The best fitting specification is chosen according to the Likelihood Ratio (LR) results and the minimum value of the Information Criteria (IC) (not reported). Once heteroscedasticity in the conditional mean

Table 4.60. AR Lags

Japan	$1, 2, 3, 5, 6, 8, 9, 10$
South Korea	$1, 2, 3, 9, 10, 12$
Singapore	$1, 2, 3, 4, 5, 8, 12$
Taiwan	$1, 3, 4, 6, 10, 12$

Notes: The numbers represent the AR lags used in the mean equations of the model.

has been accounted for, an AR(12) specification appears to capture the serial correlation in all four inflation series. Table 4.60 reports AR lags for the conditional means of inflation series.

The existence of outliers causes the distribution of inflation to exhibit excess kurtosis. To accommodate the presence of such leptokurtosis, one should estimate the PARCH models using non-normal distributions. As reported by Palm (1996), the use of a student-t distribution is widespread in the literature. In accordance with this, we estimate all the models using two alternative distributions: the normal and the student-t. Moreover, we allow for the possibility of structure break in the inflation data. The mean equation is modified to include structure break dummy variables. In other words, the dummy variables are included to reflect the difference between high inflation periods (oil shock, growth strategy) and low/normal inflation period in the series. We find that dummies in all four countries/regions are statistically significant.

For all countries/regions, we find the leverage term ζ to be insignificant and therefore we re-estimate the model excluding this parameter. The estimated α and β parameter are highly significant in all cases. In order to distinguish the general PARCH model from a version in which is fixed to a specific value we will hereafter refer to the latter as (P)ARCH.

Table 4.61 reports the (P)ARCH models with normal distribution applied on the innovations term. The IC chooses a (P)ARCH model with estimated power term parameter $\delta = 0.5, 0.5, 0.5$ and 0.6 for Japan, South Korea, Singapore and Taiwan, respectively.

Table 4.61. (P)ARCH Models

	Japan	South Korea	Singapore	Taiwan
		Normal		
α	0.158	0.138	0.279	0.332
	(4.124)	(3.733)	(4.980)	(3.736)
β	0.809	0.869	0.674	0.565
	(17.030)	(21.426)	(7.792)	(4.680)
δ	0.5	0.5	0.5	0.6

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the (P)ARCH models. The numbers in parentheses are z-Statistic. r are the degrees of freedom of Student-t distribution.

Next, we report the estimation results of an AR-(P)ARCH-M model of inflation, with $g(h_{t)} = h_t$, for all four countries/regions. Table 4.62 reports only the estimated parameters of interest. In Japan, South Korea and Singapore the estimates for the "in-mean" parameter (k) are statistically significant (see the "Mean" panel of Table 4.62). In these three countries, there is evidence in favour of the Cukierman-Meltzer hypothesis as the value of the "in-mean" coefficient is positive. We also find negative "in-mean" effect for Taiwan which is in accordance with the Holland hypothesis. For Japan, South Korea and Singapore, the values of the "power" coefficients are 0.6. For Taiwan, the power term is 1.0.

In what follows, we report the estimation results of an AR-PARCH-L model of inflation in the four countries/regions with lagged inflation included in the conditional variance as the "level effect". In the expressions for the conditional variances reported in Table 4.62, various lags of inflation (from 1 to 12) were considered with the best model chosen on the basis of the minimum value of the AIC (see the "Level" panel of Table 4.62). For all countries/regions, there is strong evidence that inflation affects its uncertainty positively as predicted by Friedman (1977) and Ball (1992). The estimated (absolute) "level" coefficient is in the range $0.036 < |\gamma_i| < 0.086$. The chosen value of δ are 0.5 (Japan, Singapore) and 0.6 (South Korea, Taiwan).

Finally, Table 4.62 also reports the estimation results of an AR-(P)ARCH-ML model. That is, we estimate a system of equations that allows not only the current value of the conditional variance to affect average inflation but also allows lag of the latter to influence the former (see the “Mean-Level” panel of Table 4.62). All estimated “level” coefficients are highly significant. As with the L model, we again find support for Friedman’s hypothesis in all four countries/regions. The “level” parameter is in the range $0.039 < |\gamma_i| < 0.206$. Moreover, we show evidence regarding the direction of the impact of a change in nominal uncertainty on inflation. That is, we find evidence in favour of the Cukierman-Meltzer hypothesis in Japan, South Korea and Singapore. The “in-mean” parameter is negative but insignificant in Taiwan. Japan is the country with the highest “risk premium” parameter (0.568). As with the M models in all four countries/regions, the values of the “power” coefficients are 0.5 except Taiwan. When we include “level” effects the impact of uncertainty on inflation holds in the similar level. On the other hand, the impact of inflation on its uncertainty is robust to the inclusion or exclusion of “in-mean” effects.

Table 4.63 reports for Japan estimates of the k parameters of the (P)ARCH-M model with $g(h_t) = h_t$ and errors that are conditionally normal, for various positive δ. The estimated values of the “in mean” effect are sensitive to changes in the ”power” term. Note that the statistical significance of the “risk premium” decreases as the value of δ increases (see t-statistic in brackets in Table 4.63). The most interesting finding is that the autocorrelation function of $|\pi_t|^d$ (for lag 12) reaches its maximum, approximately, at this point.

4.5 Robustness Check

Table 4.62. (P)ARCH-ML Models, Normal Distribution

	Japan	South Korea	Singapore	Taiwan
	Mean			
k	0.928	0.117	0.210	-0.087
	(2.778)	(1.879)	(2.672)	(1.721)
δ	0.6	0.6	0.6	$1.0^{\dagger}$
	Level			
γ_i	0.082	0.036	0.074	0.086
	(3.154)	(2.079)	(1.614)	(4.417)
	{1}	{1}	{1}	{1}
δ	0.5	0.6	0.5	0.6
	Mean-level			
k	0.568	0.158	0.282	-0.037
	(1.796)	(2.638)	(2.748)	(0.699)
γ_i	0.071	0.039	0.066	0.206
	(3.045)	(2.824)	(1.775)	(3.652)
	{1}	{1}	{1}	{1}
δ	0.5	0.5	0.5	$1.4^{\ddagger}$

Notes: For each of the five Asian countries/regions, this table reports estimates of the parameters of interest for the various (P)ARCH-M, (P)ARCH-L and (P)ARCH-ML models.
In all cases $g(h_t) = h_t$. † No convergence with $\delta < 1.0$.
‡ No convergence with $\delta < 1.4$.
The numbers in parentheses are z-Statistic. r are the degrees of freedom of Student's t distribution. The numbers in { } indicate the lags of the 'level' terms.

Table 4.63. (P)ARCH-M Models for Japan, Normal Distribution

δ	0.5	0.6	0.8	1.0	1.2	1.5	1.8	2.0	2.5
k	0.813	0.928	0.406	0.291	0.409	0.357	0.311	0.302	0.306
	(2.719)	(2.778)	(1.572)	(1.094)	(1.516)	(1.333)	(1.208)	(1.195)	(1.262)
AIC	0.768	**0.760**	0.770	0.776	0.777	0.783	0.788	0.792	0.799
LL	-192.54	-190.15	-192.81	-194.51	-194.88	-196.43	-197.83	-198.71	-200.68

Notes: This table reports estimates of the "in mean" parameters for the (P)ARCH-M models.
In all cases $g(h_t) = h_t$. The numbers in parentheses are z-Statistic.
The bold numbers indicate the minimum value of the AIC. LL denotes the maximum log-likelihood value.

Table 4.64. (P)ARCH models with Structural Break

	Japan	South Korea	Singapore	Taiwan
α	0.207 (4.492)	0.178 (4.532)	0.255 (5.698)	0.404 (3.482)
β	0.715 (10.201)	0.838 (21.056)	0.691 (10.862)	0.411 (2.933)
D_i	-0.217 (5.497)	-0.710 (8.438)	-0.110 (2.737)	-0.417 (4.557)
δ	0.5	0.6	0.5	0.8

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the (P)ARCH models with structural break.
The numbers in parentheses are z-Statistic.
r are the degrees of freedom of Student-t distribution.

4.5.1 Structural Break

In order to capture the difference between high inflation periods (oil shock, export-oriented growth strategy) and low/normal inflation period in the series, we use the methodology developed by Bai and Perron (2003) to examine whether there is any structural break in inflation and when it occurred. Bai and Perron (2003) address the problem of testing for multiple structural changes under very general conditions on the data and the errors. In addition to testing for the existence of breaks, these statistics identify the number and location of multiple breaks.

Following Bai-Perron methodology, we detected structural break points at February of 1981 for Japan, October of 1981 for South Korea, March of 1982 for Singapore and August of 1981 for Taiwan, respectively.

Table 4.64 reports the (P)ARCH models with structural break dummy variable incorporated in the mean equation and with normal distribution applied on the innovations term. The IC chooses a (P)ARCH model with estimated power term parameter of $\delta = 0.5, 0.6, 0.5$ and 0.8 for Japan, South Korea, Singapore and Taiwan respectively. In all cases, the coefficients for the structural break dummy variables are negative and significant.

Next, to check the sensitivity of our results to the presence of structural break in the inflation data, we report the estimation results of an AR-(P)ARCH-M model of inflation with structural break dummy variable incorporated in the mean equation and $g(h_t) = h_t$ for all four countries/regions. Table 4.65 reports only the estimated parameters of interest and in all cases, the coefficients for the structural break dummy variables are negative and significant. Regarding to "level effect", the results are very similar to those obtained without using dummy variables, all four countries/regions' inflation have positive impact on its uncertainty, the strong evidence in support of the Friedman hypothesis in all countries/regions is invariant to the inclusion or exclusion of the "in-mean" effect. Regarding the reverse causal effect, our evidence is country/region specific. As before, the estimates for the "in-mean" parameter (k) in Japan and Singapore are positive $(0.315$ and $0.162)$ and statistically significant at 10% levels. Very interestingly, the "in mean" effect for South Korea turns around, it become negative and significant. The impact of nominal uncertainty on inflation in Taiwan is remaining negative and significant at 2% level. So, when we take structural break into account, we find evidence in support of the Cukierman-Meltzer theory in Japan and Singapore and evidence for the Holland hypothesis in South Korea and Taiwan. When we account for "level" effects, the influence of nominal uncertainty on inflation becomes stronger in Japan and Singapore and become weaker in South Korea and Taiwan. All power term parameters of δ are below 1, except the ML model for Taiwan.

Table 4.66 reports, again for Japan for the same reasons as before, estimates of the parameters of the (P)ARCH-M model with structural break dummy and $g(h_t) = h_t$, for various positive δ. Similar to our sensitivity analysis before, the estimated values of the "in mean" effect are sensitive to changes in the "power" term. It is important to mention that when the errors are conditionally normal and we incorporate structural dummies in the model the statistical significance of the "risk premium" decreases as the value of δ increases. the AIC is minimized when $\delta = 0.5$, note that the autocorrelation function of $|\pi_t|^d$ (for lag 12) reaches its maximum at this point.

Table 4.65. (P)ARCH-ML Models with Structrual Break, Normal Distribution

	Japan	South Korea	Singapore	Taiwan
		Mean		
k	0.315 (1.833)	-0.248 (1.921)	0.162 (1.682)	-0.134 (2.556)
δ	0.5	0.9	0.7	$0.9^{\dagger}$
		Level		
γ_i	0.096 (4.140) $\{1\}$	0.106 (3.394) $\{1\}$	0.061 (1.886) $\{1\}$	0.112 (4.369) $\{1\}$
δ	0.5	0.9	0.5	0.7
		Mean-level		
k	0.427 (1.736)	-0.141 (1.788)	0.318 (3.725)	-0.088 (2.375)
γ_i	0.080 (4.055) $\{1\}$	0.093 (3.098) $\{1\}$	0.070 (4.086) $\{1\}$	0.208 (3.101) $\{1\}$
δ	0.5	0.8	0.5	$1.4^{\ddagger}$

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the various (P)ARCH-M,(P)ARCH-L and (P)ARCH-ML models with structural break.
In all cases $g(h_t) = h_t$.
† No convergence with $\delta < 0.9$.
‡ No convergence with $\delta < 1.4$.
The numbers in parentheses are z-Statistic. r are the degrees of freedom of Student's t distribution. The numbers in { } indicate the lags of the 'level' terms.

Table 4.66. (P)ARCH-M Models with Structrual Break for Japan, Normal Distribution

δ	0.4	0.5	0.6	0.8	1.0	1.5	1.8	2.0	2.5
k	0.361 (2.094)	0.315 (1.833)	0.257 (1.124)	0.211 (0.996)	0.111 (0.472)	0.157 (0.641)	0.160 (0.666)	0.174 (0.746)	0.206 (0.924)
AIC	0.796	**0.745**	0.747	0.749	0.754	0.762	0.765	0.770	0.777
LL	-198.91	-185.21	-185.70	-186.28	-187.52	-189.66	-197.83	-191.90	-193.87

Notes: This table reports estimates of the "in mean" parameters for the (P)ARCH-M models.
In all cases $g(h_t) = h_t$. The numbers in parentheses are z-Statistic.
The bold numbers indicate the minimum value of the AIC. LL denotes the maximum log-likelihood value.

Table 4.67. (P)ARCH Models,Student-t Distribution

	Japan	South Korea	Singapore	Taiwan
	without structural dummy			
α	0.145 (4.311)	0.116 (4.387)	0.339 (5.956)	0.266 (4.867)
β	0.863 (25.291)	0.903 (43.989)	0.643 (10.511)	0.533 (5.145)
δ	0.7	0.7	0.7	0.6
r	5.593 (5.245)	5.298 (4.391)	4.637 (4.803)	3.701 (5.910)
	with structural dummy			
α	0.252 (5.029)	0.238 (4.395)	0.383 (5.951)	0.265 (4.738)
β	0.728 (12.958)	0.897 (39.801)	0.582 (8.244)	0.491 (4.050)
D_i	−0.241 (5.228)	−0.557 (6.091)	−0.049 (1.762)	−0.251 (3.698)
δ	0.7	0.7	0.7	0.6
r	5.507 (5.336)	5.736 (4.385)	4.430 (4.728)	3.553 (5.997)

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the (P)ARCH models. The numbers in parentheses are z-Statistic. r are the degrees of freedom of Student-t distribution.

4.5.2 Student-t Distribution

One important robustness test regards the role of error distribution. To check the sensitivity of our results to the distribution of the innovations we are also using the student-t distribution. Table 4.67 reports the (P)ARCH models with student-t distribution applied in the error term. Compare with corresponding models with normal distribution, IC prefer higher values of δ.

In general, the results are very similar to those obtained when the innovations are drawn from the normal distribution (see Table 4.68 and 4.69). That is, in all four countries/regions, inflation has a positive impact on its uncertainty. Regarding the reverse causal effect, our evidence is country specific. In particular, in the models without structure break dummy, we find inflation uncertainty positively impact inflation in Japan, South Korea and Singapore. When structure break is considered, the influence from

Table 4.68. (P)ARCH-ML Models, Student-t Distribution

	Japan	South Korea	Singapore	Taiwan
		Mean		
k	1.272	0.199	0.144	−0.014
	(4.049)	(2.768)	(2.798)	(0.276)
δ	0.7	0.8	0.5	$1.4^{\dagger}$
r	4.998	4.957	4.859	3.898
	(5.354)	(4.608)	(4.596)	(5.775)
		Level		
γ_i	0.093	0.040	0.082	0.107
	(2.923)	(2.208)	(2.666)	(4.339)
	{1}	{1}	{1}	{1}
δ	0.7	0.8	0.6	0.6
r	5.911	5.666	4.999	4.602
	(5.040)	(4.162)	(4.947)	(5.983)
		Mean-level		
k	0.439	0.275	0.136	0.0002
	(1.598)	(2.860)	(1.764)	(0.004)
γ_i	0.091	0.044	0.080	0.230
	(2.935)	(2.569)	(2.655)	(3.329)
	{1}	{1}	{1}	
δ	0.8	0.8	0.7	$1.4^{\dagger}$
r	6.045	5.231	4.833	4.382
	(5.112)	(4.402)	(4.890)	(5.982)

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the various (P)ARCH-M, (P)ARCH-L and (P)ARCH-ML models with structural break and assuming Student's t distribution for the error term. In all cases $g(h_t) = h_t$.

† No convergence with $\delta < 1.4$.

The numbers in parentheses are z-Statistic.

r are the degrees of freedom of Student's t distribution.

The numbers in { } indicate the lags of the 'level' terms.

inflation uncertainty on inflation is positive for Japan and Singapore while, negative for South Korea and Taiwan (insignificant).

4.5.3 In-mean Structure

Next, to check the sensitivity of our results to the form in which the time varying variance enters the specification of the mean, we also use either the conditional standard deviation or the logarithm of the conditional variance as regressor in the mean equation (see Table 4.70 4.71 4.72 and 4.73). The strong evidence in support of the Friedman hypothesis in all countries/regions is invariant to the different form of the "in-mean" effect and the "in-

Table 4.69. (P)ARCH-ML Models with Structrual Break, Student-t Distribution

	Japan	South Korea	Singapore	Taiwan
	Mean			
k	0.471 (2.159)	−0.050 (0.429)	0.133 (2.005)	−0.052 (0.923)
δ	0.8	0.7	0.6	$1.4^{\dagger}$
r	4.173 (5.623)	5.875 (4.450)	4.424 (8.436)	3.859 (5.678)
	Level			
γ_i	0.083 (2.941) {1}	0.053 (3.616) {1}	0.088 (3.139) {1}	0.089 (4.583) {1}
δ	0.5	0.5	0.5	0.5
r	6.595 (5.037)	7.471 (3.202)	4.447 (4.729)	4.639 (6.114)
	Mean-level			
k	0.230 (1.711)	−0.112 (1.729)	0.115 (1.815)	−0.040 (0.574)
γ_i	0.072 (2.847) {1}	0.048 (4.142) {1}	0.093 (3.275) {1}	0.222 (3.332)
δ	0.5	0.5	0.5	$1.4^{\dagger}$
r	6.307 (5.352)	7.818 (3.072)	4.356 (4.787)	4.421 (5.929)

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the various (P)ARCH-M, (P)ARCH-L and (P)ARCH-ML models with structural break and assuming Student's t distribution for the error term. In all cases $g(h_t) = h_t$.

† No convergence with $\delta < 1.4$.

The numbers in parentheses are z-Statistic.

r are the degrees of freedom of Student's t distribution.

The numbers in { } indicate the lags of the 'level' terms.

Table 4.70. (P)ARCH-ML Models,Varying Variance in Mean, Normal Distributions

	$g(h_t)=\sqrt{h_t}$				$g(h_t)=ln(h_t)$			
	Japan	South Korea	Singapore	Taiwan	Japan	South Korea	Singapore	Taiwan
	Mean							
k	0.531 (2.616)	0.377 (4.688)	0.224 (3.853)	−0.048 (0.397)	0.044 (2.125)	0.050 (2.648)	0.055 (2.334)	−0.037 (0.699)
δ	0.5	0.5	0.6	0.8	0.5	0.5	0.6	0.5
	Mean-level							
k	0.547 (2.337)	0.471 (3.423)	0.194 (1.842)	0.109 (0.981)	0.096 (2.343)	0.186 (4.665)	0.056 (2.467)	0.033 (0.356)
γ_i	0.062 (3.349) {1}	0.040 (2.540) {1}	0.072 (1.749) {1}	0.107 (4.668) {1}	0.067 (3.643) {1}	0.098 (3.795) {1}	0.052 (1.652) {1}	0.097 (5.060) {1}
δ	0.5	0.6	0.5	0.7	0.5	0.5	0.6	0.6

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the various(P)ARCH-M and (P)ARCH-ML models and assuming normal distribution for the error term.
The numbers in parentheses are z-Statistic.
The numbers in {} indicate the lags of the ”level” terms.

mean” effects are the same as our previous results in the models without structural break. In the models with structural break incorporated, the signs of all “in-mean” coefficients are same with the results form models which with conditional variance in mean, however, these effects become insignificant in Japan and South Korea when error term assumed to be normal distributed, and none of them are statistically significant when student-t distribution applied in the error. This result indicates that the models with conditional variance as measurement of uncertainty have most robust results.

4.6 Conclusions

The results presented above carry noteworthy implications for macroeconomic modelling and policy making. Our very strong evidence on the Friedman hypothesis is in broad agreement with the findings of the overwhelming majority of empirical studies. The country-specific evidence on the Cukierman-Meltzer hypothesis is anticipated given that national central banks adjust their rate of money growth differently to nominal uncer-

Table 4.71. (P)ARCH-ML Models wih Structrual Break,Varying Variance in Mean, Normal Distributions

	$g(h_t) = \sqrt{h_t}$				$g(h_t) = ln(h_t)$			
	Japan	South Korea	Singapore	Taiwan	Japan	South Korea	Singapore	Taiwan
				Mean				
k	0.027	−0.173	0.237	−0.347	0.002	−0.025	0.062	−0.178
	(0.161)	(1.483)	(3.023)	(1.830)	(0.129)	(0.653)	(3.604)	(2.004)
δ	0.7	0.5	0.7	0.8	0.5	0.8	0.5	0.6
				Mean-level				
k	0.203	−0.084	0.215	−0.366	0.091	−0.007	0.039	−0.156
	(0.660)	(1.060)	(2.559)	(1.747)	(0.220)	(0.234)	(2.495)	(1.723)
γ_i	0.105	0.058	0.059	0.126	0.112	0.071	0.065	0.176
	(4.417)	(3.037)	(1.888)	(3.897)	(4.360)	(3.125)	(1.865)	(3.171)
	{1}	{1}	{1}	{1}	{1}	{1}	{1}	{1}
δ	0.6	0.5	0.5	0.8	0.7	0.6	0.6	$1.2^{\dagger}$

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the various (P)ARCH-M and (P)ARCH-ML models with structural break and assuming normal distribution for the error term.

†No convergence with $\delta < 1.2$.

The numbers in parentheses are z-Statistic.

The numbers in {} indicate the lags of the ”level” terms.

Table 4.72. (P)ARCH-ML Models,Varying Variance in Mean, Student-t Distributions

	$g(h_t) = \sqrt{h_t}$				$g(h_t) = ln(h_t)$			
	Japan	South Korea	Singapore	Taiwan	Japan	South Korea	Singapore	Taiwan
				Mean				
k	0.695	0.300	0.054	0.217	0.085	0.055	0.030	0.072
	(3.350)	(2.934)	(14.251)	(1.605)	(2.687)	(2.043)	(1.981)	(1.074)
δ	0.8	0.7	0.7	0.6	0.9	0.8	0.7	0.6
r	4.897	5.140	4.637	3.490	4.853	5.225	4.847	3.618
	(84.015)	(4.478)	(4.803)	(5.999)	(5.313)	(4.432)	(4.671)	(5.917)
				Mean-level				
k	0.557	0.388	0.182	0.231	0.074	0.343	0.036	0.072
	(1.764)	(2.751)	(1.943)	(1.294)	(1.868)	(6.412)	(1.855)	(1.118)
γ_i	0.097	0.043	0.089	0.103	0.105	0.177	0.090	0.106
	(2.982)	(2.668)	(2.807)	(4.402)	(3.084)	(5.539)	(2.864)	(4.185)
	{1}	{1}	{1}	{1}	{1}	{1}	{1}	{1}
δ	0.8	0.8	0.7	0.6	0.7	0.9	0.7	0.6
r	5.991	5.367	4.893	4.487	5.976	3.201	4.913	4.620
	(5.122)	(4.219)	(4.852)	(5.989)	(5.199)	(5.747)	(4.816)	(5.933)

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the various (P)ARCH-M and (P)ARCH-ML models with structural break and assuming Student’s t distribution for the error term. The numbers in parentheses are z-Statistic. r are the degrees of freedom of Student’s t distribution. The numbers in {} indicate the lags of the ”level” terms.

Table 4.73. (P)ARCH-ML Models with Structrual Break,Varying Variance in Mean, Student-t Distributions

	$g(h_t) = \sqrt{h_t}$				$g(h_t) = ln(h_t)$			
	Japan	South Korea	Singapore	Taiwan	Japan	South Korea	Singapore	Taiwan
	Mean							
k	0.136	−0.074	0.043	−0.165	0.027	−0.009	0.019	−0.074
	(1.107)	(0.457)	(0.457)	(0.959)	(0.129)	(0.263)	(0.999)	(0.930)
δ	0.6	0.9	0.7	0.8	0.7	0.9	0.7	0.6
r	5.502	5.840	4.492	3.741	5.201	5.827	4.617	3.691
	(5.357)	(4.479)	(4.702)	(5.801)	(5.249)	(4.485)	(10.168)	(5.987)
	Mean-level							
k	0.292	−0.076	0.105	−0.058	0.027	−0.008	0.025	−0.042
	(1.308)	(0.477)	(1.045)	(0.289)	(1.027)	(0.261)	(1.142)	(0.473)
γ_i	0.098	0.083	0.100	0.141	0.112	0.057	0.105	0.109
	(3.059)	(3.323)	(2.924)	(4.056)	(3.032)	(3.120)	(3.017)	(4.416)
	{1}	{1}	{1}	{1}	{1}	{1}	{1}	{1}
δ	0.6	0.9	0.7	0.8	0.7	0.7	0.7	0.6
r	6.211	6.908	4.682	4.571	6.241	6.984	4.675	4.718
	(5.122)	(3.310)	(4.798)	(6.055)	(5.470)	(3.392)	(4.778)	(6.144)

Notes: For each of the four Asian countries/regions, this table reports estimates of the parameters of interest for the various (P)ARCH-M and (P)ARCH-ML models with structural break and assuming Student's t distribution for the error term. The numbers in parentheses are z-Statistic. r are the degrees of freedom of Student's t distribution. The numbers in {} indicate the lags of the "level" terms.

tainty depending on their relative preference towards inflation stabilisation. Previous literature reports mixed results that are sensitive to factors such as the measure of uncertainty and the countries examined. In general, when we use the value of the "power" term that is preferred by the IC, we find that the evidence in support of the Cukierman-Meltzer hypothesis for Japan and Singapore are robust to i) the distribution of the innovations and ii) the presence of structural break in the inflation data. We show, however, that the significance of the "in-mean" effect is sensitive to the choice of the "heteroscedasticity" parameter. Moreover, we find positive "in-mean" effect for South Korea in the models without structural dummy and negative "in mean" effect when the structural dummy is incorporated. Though, we find negative "in mean" effect for Taiwan, and this effect only became statistically significant when the model be estimated with normal distribution applied on the error term and structure break was considered. One possible reason for these differences among countries/regions is that they follow different monetary policies and dispose of different Central Banking institutions.

Our results on Japan and Singapore are highly consist with other empirical literatures, such as, Grier and Perry (1998), Conrad and Karanasos (2005b), Chen et al (2008) and Jiranyakul and Opiela (2010). These results are well explained by existing theories, given that Japan and Singapore are generally been considered as exchange rate targeting countries, rather than monetary authorities which adopt inflation rate targeting policy. Referring to Taiwan, the empirical evidence for the effect of uncertainty on inflation is unclear, for example, Chen et al (2006) use moving average standard deviation as proxy of uncertainty find positive linear effect and negative nonlinear effect on inflation. Chen et al (2008) claim a invented U-shape relationship when they describe the impact of inflation uncertainty on inflation itself. This may due to the "growth under the stable environment" policy that Taiwan monetary authority are pursuing, that is, promote economy when price stability within certain threshold (Shen and Hakes, 1995). Shouth Korea seems to be a good evidence of Cukierman-Meltzer hypothesis as South Korean authority adopted "export-oriented growth strategy" in 1960s and 1970s and they controlled exchange rate tightly (Start to transit to floating system at the end of 1997). Thomton(1997) and Chen et al (2008) both find very significant results in favour of Cukierman-Meltzer hypothesis in South Korea with data that covers different growth strategy periods. However, none of them took this possible structure break in to consideration. Our result of South Korea in line with these two researches when we estimate our models without structural dummy variable, but when structural dummy is incorporated in, the result reveal a sharp contrast: inflation uncertainty impact on inflation negatively and significantly, no matter the error term draw from normal distribution or student-t distribution. This dramatic outcome in favour of Holland hypothesis chime in with Dueker and Kim (1999) who claim that inflation was strictly controlled by South Korean monetary authority even in the "export-oriented growth strategy" period, the upper threshold of tolerable annual inflation for the Bank of Korea was about 20% before 1983 and 6% after.

Our results also highlight the importance of using the PARCH specification to model the power transformation of the conditional variance of inflation. It increases the flexibility of the conditional variance specification by allowing the data to determine the power of inflation, for which the predictable structure in the volatility pattern is the strongest. The statistical significance of the "in-mean" effect is highly dependent on the choice of the value of the "heteroscedasticity" parameter. For normal distribution, the effect becomes insignificant if the "power term" surpasses a specific value. This suggests that if we had assumed a prior linear relationship between inflation and its uncertainty, the so-called Bollerslev specification, we would not detect any significant link between the two variables. Most interestingly, this value coincides with the one chosen by the IC and the one for which the sample autocorrelation of the power-transformed inflation series is maximal. Whether this coincidence is systematic will be the focus of further research.

4.A Appendix

Table 4.74. Power Term Parameters of (P)ARCH-ML Models

	Normal Distribution				Student-t Distribution			
	Japan	South Korea	Singapore	Taiwan	Japan	South Korea	Singapore	Taiwan
	Without Structural Break							
Simple	0.5	0.5	0.5	0.6	0.7	0.7	0.7	0.6
Level	0.5	0.6	0.5	0.6	0.7	0.8	0.6	0.6
	With Structural Break							
Simple	0.5	0.6	0.5	0.8	0.7	0.7	0.7	0.6
Level	0.5	0.9	0.5	0.7	0.5	0.5	0.5	0.5

Notes: For each of the four Asian countries/regions, this table reports the power term Parameters of the various (P)ARCH and (P)ARCH-L models.

Table 4.75. Power Term Parameters of (P)ARCH-ML Models with Conditional Variance in Mean

	Normal Distribution				Student-t Distribution			
	Japan	South Korea	Singapore	Taiwan	Japan	South Korea	Singapore	Taiwan
	Without Structural Break							
Mean	0.6	0.6	0.6	$1.0^{\dagger}$	0.7	0.8	0.5	$1.4^{\ddagger}$
Mean-level	0.5	0.5	0.5	$1.4^{\ddagger}$	0.8	0.8	0.7	$1.4^{\ddagger}$
	With Structural Break							
Mean	0.5	0.9	0.7	$0.9^{\star}$	0.8	0.7	0.6	$1.4^{\ddagger}$
Mean-level	0.5	0.8	0.5	$1.4^{\ddagger}$	0.5	0.5	0.5	$1.4^{\ddagger}$

Notes: For each of the four Asian countries/regions, this table reports the power term Parameters of the various (P)ARCH-M and (P)ARCH-ML models. In all cases $g(h_t) = h_t$.

†No convergence with $\delta < 1.0$

*No convergence with $\delta < 0.9$

‡No convergence with $\delta < 1.4$

Table 4.76. Power Term Parameters of (P)ARCH-ML Models with Conditional Standard Deviation in Mean

	Normal Distribution				Student-t Distribution			
	Japan	South Korea	Singapore	Taiwan	Japan	South Korea	Singapore	Taiwan
Without Structural Break								
Mean	0.5	0.5	0.6	0.8	0.8	0.7	0.7	0.6
Mean-level	0.5	0.6	0.5	0.7	0.8	0.8	0.7	0.6
With Structural Break								
Mean	0.7	0.5	0.7	0.8	0.6	0.9	0.7	0.8
Mean-level	0.6	0.5	0.5	0.8	0.6	0.9	0.7	0.8

Notes: For each of the four Asian countries/regions, this table reports the power term Parameters of the various (P)ARCH-M and (P)ARCH-ML models. In all cases $g(h_t) = \sqrt{h_t}$.

Table 4.77. Power Term Parameters of (P)ARCH-ML Models with Logarithm of the Conditional Variance in Mean

	Normal Distribution				Student-t Distribution			
	Japan	South Korea	Singapore	Taiwan	Japan	South Korea	Singapore	Taiwan
Without Structural Break								
Mean	0.5	0.5	0.6	0.5	0.9	0.8	0.7	0.6
Mean-level	0.5	0.5	0.6	0.6	0.7	0.9	0.7	0.6
With Structural Break								
Mean	0.5	0.8	0.5	0.6	0.7	0.9	0.7	0.6
Mean-level	0.7	0.6	0.6	$1.2^{\dagger}$	0.7	0.7	0.7	0.6

Notes: For each of the four Asian countries/regions, this table reports the power term Parameters of the various (P)ARCH-M and (P)ARCH-ML models. In all cases $g(h_t) = ln(h_t)$.

†No convergence with $\delta < 1.2$

Chapter 5
Dual Long-memory and the Link between Bid-ask Spread, Volatility and Volume in FTSE100

5.1 Introduction

One of the most critical factors that investors look for in any market is liquidity. Liquidity is defined as the ability to trade stock rapidly with little price impact. To maintain liquidity, exchanges use market makers, who are individuals willing to provide a financial market whenever the investors wish to trade. In return for providing the financial market, market makers are granted monopoly rights by the exchange to post different prices for stock purchases and sales. As a result, market makers buy stock at the bid price and sell the stock at the higher ask price. This ability to buy the stock low and sell high is the market makers' compensation for providing liquidity in the financial market. Their compensation is defined as the ask price minus the bid price, which in turn is defined as the bid-ask spread.

Therefore the ask (bid) price is the price that investors actually pay to purchase (sell) stock. Investors pay these prices instead of the quoted stock price to ensure that market makers are compensated for providing liquidity in the financial market[71]. Purchases and sales of assets at the ask and bid prices are common in financial markets. For example, derivatives asset pricing of both call and put options is found to be more closely related to the ask and bid prices rather than the usually employed mid-price, Gregoriou et al (2007).

[71] Note the quoted stock price is usually the mid price between the ask and bid prices. The mid price is defined as ((ask price – bid price)/2).

Studies such as Atkins and Dyl (1997), Constantinides (1986), Glosten and Harris (1988), Menyah and Paudyal (2000) and Stoll (1989) relate the spread or the change in the spread to a vector of characteristics that are associated with the individual securities. These factors identified in prior spread models are the market value of the firm, the risk of the security (usually approximated by the standard deviation of returns), and trading volume.

This chapter tries to contribute to the study of the relationship between spread, volatility, and trading volume in FTSE100 stock index by investigating two-way causality relationships using a bivariate AR-FI-GARCH model and multiple measurements of risk and spread. Specially, we distinguish the inventory cost component of the bid-ask spread and the information cost component of the bid-ask spread. the impact of electronic trading system also be considered.

5.2 Literature Review

Early research on the relationship between bid-ask spread and volatility includes Benston and Hagerman (1974), Stoll (1978) and Glosten and Milgrom (1985) who find that price volatility is one of the key explanatory variables on bid-ask spread and have positive effects on bid-ask spread. Roll (1984) and Karpoff (1986) argue that market makers widen the bid-ask spread when the variance of returns increases in order to compensate the cost with increasing risk in security market, this is because the higher the risk, the harder it is for the market maker to trade the stock, and the market maker has to increase the bid-ask spread in order to compensate the great cost of inventory. Admati and Pfleiderer (1988) and Kim and Verrecchia (1994) also point out that dealers widen the bid-ask spread to compensate for their informational disadvantage when informed trading increase which is implied by the rise in the risk. Informed traders tend to be more active when liquidity trading is concentrated, since price is more informative in these periods and this interaction between informed traders and liquidity traders leads to higher volatility of stock price (Admati and Pfleiderer, 1988). The positive relationship between bid-ask spread and volatility in various markets has been proved by many studies such as Ding and Chong (1997) who study the Nikkei stock index futures trading in the Singapore Monetary Exchange (SIMEX). Their results implied a positive correlation between bid-ask spread and volatility. Wang and Yau (2000) use a GMM estimation procedure investigate the S&P500, the deutsche mark, silver and gold futures contracts, the results show that price volatility and bid-ask spread are positively related, and they also find negative relationship between trading volume and bid-ask spread. Ding and Charoenwong (2003) show evidence of a positive relationship between volatility and bid-ask spread and a negative relationship between trading volume and bid-ask spread also exist in thinly traded equity index futures contracts on the Singapore Exchange. In agricultural futures markets, Bryant and Haigh (2004) find that for LIFFE coffee, bid-ask spread get wider when volatility rises, again bid-ask spread and volume are negatively related, but for cocoa, these relationships only appear after electronic trading was adopted. They also pointed out that electronic trading widens the bid-ask spread. Frank and Garcia (2009) suggest that in agricultural futures markets volume and volatility are endogenous variables and significantly relate to the bid-ask spread. Volatility has positive effect on the bid-ask spread while volume negatively correlated with the bid-ask spread, but their result indicate a decrease in bid-ask spread after trading was computerized, which is in line with Pirrong (1996) and Gilbert and Rijken (2006). In the stock market, Gillemot et al (2005) who investigate stocks with high capitalization in both NYSE and LSE, detect positive correlation between spread and volatility. Kanagaretnam et al (2005) study NYSE transactions around quarterly earning announcements, their results indicate a positive correlation between abnormal spread and revision volatility. Gregoriou et al (2005) also find that volatility of returns has a positive impact on the spread of FTSE 100 companies, while the same result is also found in the Australian Stock Exchange by Frino and Jones (2005). In the foreign exchange markets, the relationship between the magnitude of spread and the underlying exchange rate volatility is tested by Bollerslev and Melvin (1994) using an MA(l)-GARCH(l, 1) model. Strong positive relationship between volatility and spreads are highly statistically significant. Interestingly, Booth and Gurun (2008) following the

Bollerslev and Melvin (1994) approach find that properties of financial markets that existed 600 years ago can also be described by modern economic models and reveal same character as today's financial markets.

On the other hand, the impact of the trading cost and the information asymmetry on volatility is also gained many scholars' interests. The effect of the trading cost on volatility is still not very clear. Tobin (1978, 1984), Stiglitz (1989) and Summers and Summers (1989) argue that higher transaction costs will lead to a lower market volatility since it would discourage short-term speculation which is believed as turmoil in the market. However, opposing views include Friedman (1953), Miller (1991), Dooley (1996) who believe that short-term speculation would benefit the market stabilization as well. Grundfest and Shoven (1991) and Kupiec (1996) also point out that higher transaction cost leads to lower transaction frequency which requires larger price movement. Empirical researches show mixed results, Roll (1989) investigates 23 markets and finds negative but insignificant correlation between transaction tax and stock volatility. Mulherin (1990) gets the same result between the NYSE trading cost and the Dow Jones returns. Liu and Zhu (2009) follow the same methods of Jones and Seguin (1997) find that lower commission rate lead to higher market volatility in Japan. Norden (2009) studies the OMX Nordic Exchange before and after the reduction of the exchange fee for trading the OMXS 30 index futures with more than 22% in 2006, the result indicates that the reduction of trading cost leads to higher liquidity (higher trading volume, lower bid-ask spread) in the market, it also brings greater risk (higher volatility). Umlauf (1993) studies Swedish equity returns in the 1980s and shows positive relationship between the transaction tax and the market volatility. Positive correlation also be found in the US market by Jones and Seguin (1997), in the UK market by Saporta and Kan (1997) and in French market by Hau (2006). Green et al (2000, p.595) classify different transaction cost and identify various measures of the market volatility in London Stock Exchange, point out that "increased transactions costs generally increase market volatility, probably through a thin trading effect. However, increased transactions costs tend to reduce fundamental volatility."

French and Roll (1986) and Jones et al. (1994) point out that the private information affects price via trading. Crouzille et al. (2004) use event study on the effect of asymmetric information on the bank stock price volatility, their results show that unexpected bank stock volatility can not be easily detected by using proxies of the information asymmetry. Puffer (1991) points out that the private information in Tokyo market not only influences return volatility in Tokyo market but also affects that of New York market. McGroarty et al. (2009) follow the idea of Easley and O'Hara (1992) that the private information associates with unexpected volume and use unexpected order flow and unexpected volume as proxies for the private information, their results show that the informed trading generates volatility in the market and this volatility larger than the volatility generated by the uninformed trading.

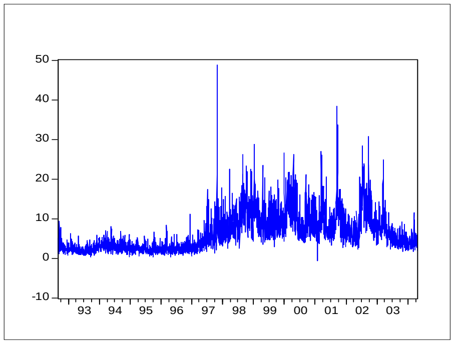
5.3 Data

Our daily data of FTSE100 index starts from 1^{st} September 1992 to 30^{th} April 2004, 2948 observations are included. Figure 5.26 shows the four measurements of spread we use, SPRD is the relative bid-ask spread, ESPRD stands for the effective bid-ask spread, CSPRD and ISPRD represent the estimated inventory cost component of the bid-ask spread and the estimated information cost component of the bid-ask spread respectively. We use two measurements of price volatility: close-to-close volatility (PVLT) and range-based volatility (GVLT) (see Figure 5.27). Figure 5.28 shows the measurement of volume: SVLM which is the 100-day backward moving average of turnover by volume.

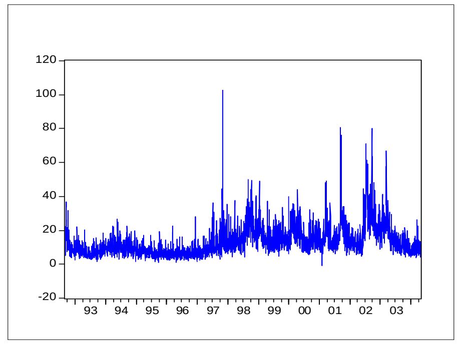
5.3.1 Spread Measures

We use four spread measures to proxy the liquidity of the stock index. The first measure is the relative bid-ask spread (SPRD) which is defined as the ask price minus the bid

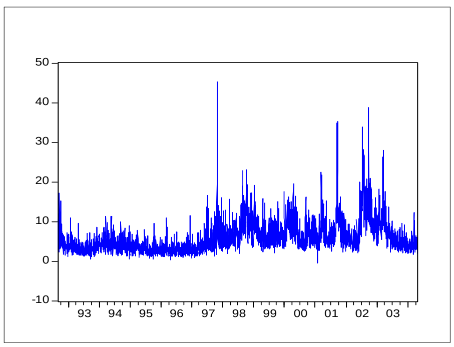
Fig. 5.26. Measurements of Spread



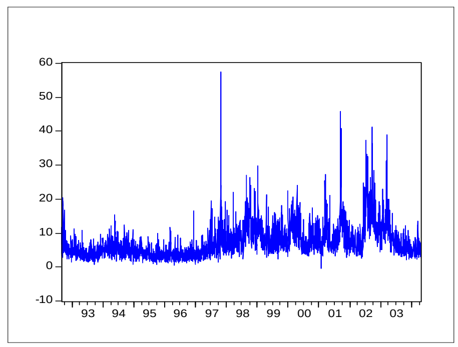
SPRD



ESPRD

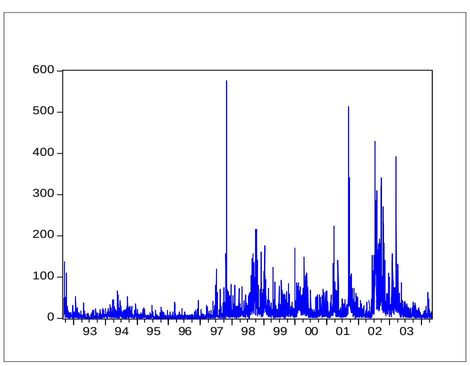


ISPRD

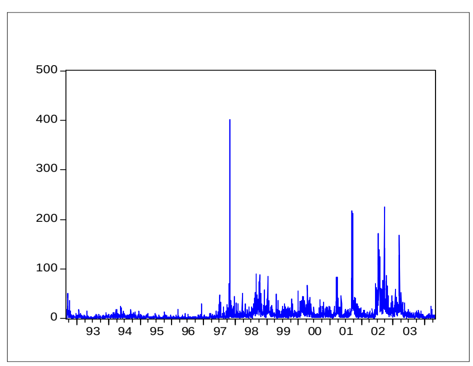


CSPRD

Fig. 5.27. Measurements of Volatility

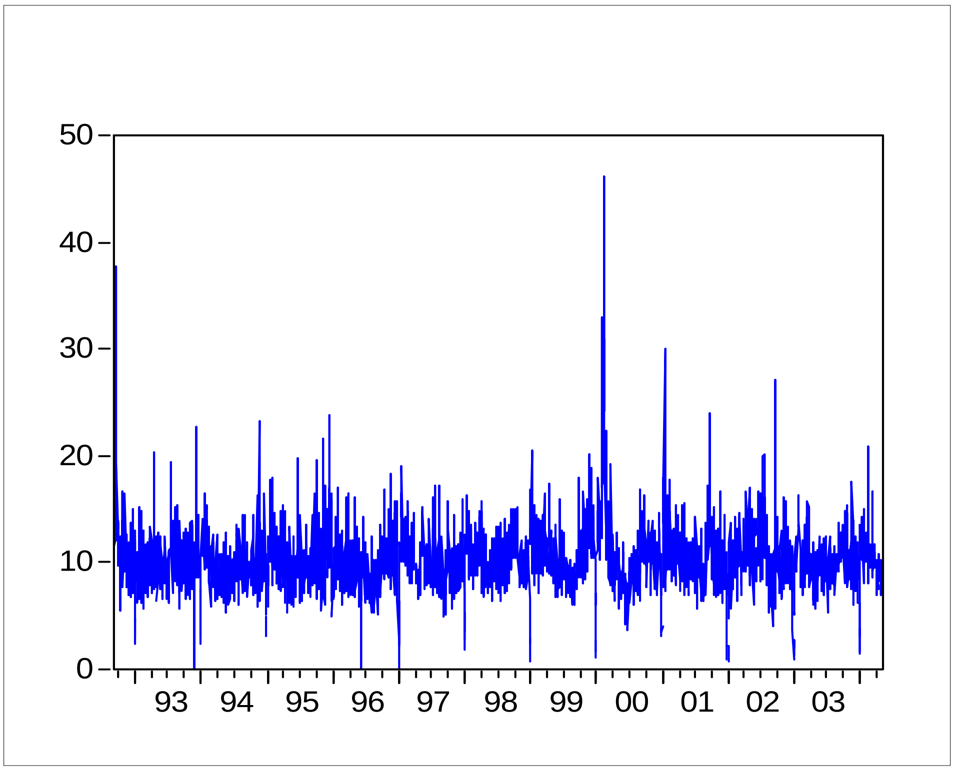


GVLT



PVLT

Fig. 5.28. Measurement of Volume



SVLM

price divided by the average of the bid and ask prices. However, Lee and Ready (1991) point out the problem with the relative spread is that it can be regarded as an inaccurate measure of liquidity, because many trades occur at prices within the bid and ask price. In order to overcome this problem we can calculate the effective bid-ask spread (ESPRD) which is the second measure of liquidity. The effective spread is measured as twice the absolute value of the difference between a transaction price and the midpoint of the bid and ask quotes at the time of transaction. Extensive theoretical literature (see among others Huang and Stoll, 1997; Lin, Sanger and Booth, 1995; Madhavan, Richardson and Roomans, 1997; to name but a few) decomposes trading costs into its non-information and information components. The non-information component comprises the direct costs of inventory holding and order processing while the information component is associated with the costs of asymmetric information. The latter is commonly known as the adverse selection costs of trading. Its isolation and use in modeling market liquidity reveals the magnitude of the influence of asymmetric information on trading costs.

We compute the adverse selection component of total trading costs following the methods of Madhavan, Richardson and Roomans (1997, henceforth MRR). The MRR propose the following model for equity price changes:

$$\Delta p_t = \alpha + (\phi + \theta)Q_t - (\phi + \rho\theta)Q_{t-1} + u_t \tag{5.24}$$

Where, Δ is the first difference operator and P_t denotes the transaction price of security at time t. Q_t is a trade initiation indicator variable which is equal to 1. The constant, α, represents the drift in prices; and u_t, a random error term, embeds the noises associated with price discreteness. ϕ measures market-makers' direct cost of supplying liquidity per share (transaction costs component). Theta (θ) is the information asymmetry parameter which measures the magnitude of the adverse selection cost. The rho (ρ) is the auto-correlation coefficient of order flow which can also be defined as $\rho = 2\gamma - (1 - \beta)$; where the parameters γ and β respectively denote the probabilities of trade flow continuation and mid-quote execution[72]. Equation (5.24) expresses changes in security price as a function of order buy flows, transaction costs, adverse selection costs and the noises associated with price discreteness. MRR suggest estimating the price formation equation by Generalized Method of Moments (GMM) under the following moment restrictions:

$$\begin{aligned} &E[Q_t Q_{t-1} - Q_t^2 \rho] = 0,\, E[|Q_{it}| - (1-\theta)] = 0,\, E[u_t - \alpha] = 0, \\ &E[(u_t - \alpha)Q_{it}] = 0,\, E[(u_t - \alpha)Q_{it-1}] = 0 \end{aligned} \tag{5.25}$$

The first moment defines the autocorrelation in trade initiation of trades, the second moment is the crossing probability, the third moment defines the drift term, α, as the average pricing error. The last two moments are OLS normal equations. We estimate the parameters of Equation (5.24) by GMM estimator, subject to the moment restrictions given in Equation (5.25), for each company of our sample. The MRR ratio of the information component r is calculated as:

$$r = \frac{\widehat{\phi}}{(\widehat{\phi} + \widehat{\theta})} \tag{5.26}$$

Thus, we calculate the information component of the spread (ISPRD) by:

$$ISPRD = r \times ESPRD$$

and the component of costs of inventory holding (CSPRD) by:

$$CSPRD = (1 - r)ESPRD$$

72 For a detailed exposition of this price evolution mechanism readers are refereed to MRR (1997).

5.3.2 Volatility Measures

We use two measurements of price volatility, The first measurement of volatility (PVLT) is the volatility estimated by classical estimation procedure that based on closing price as follows:

$$PVLT = (C_t - C_{t-1})^2 \tag{5.27}$$

where C_t is the closing price on day t. “The advantages of this classical estimator are its simplicity of usage and its freedom from obvious sources of error or bias.”(Garman and Klass, 1980, p.70).

The second measurement of volatility GVLT is a range-based volatility constructed using data on the daily high, low, opening, and closing prices in the FTSE100 index, We employ the classic range-based estimator of Garman and Klass (1980) which believed as a measurement of volatility with very little bias and is more efficient than the traditional close-to-close estimator[73]. The range-based measure of volatility (GVLT) is derived as:

$$GVLT = \frac{1}{2}u^2 - (2ln2 - 1)c^2, \quad t \in \mathbb{Z}, \tag{5.28}$$

where u and c are the differences in the natural logarithms of the high and low, and of the closing and opening prices respectively.

5.3.3 Volume Measure

Because trading volume is nonstationary, several detrending procedures for the volume data have been considered in the empirical finance literature (see, for details, Lobato and Velasco, 2000).[74] We form a trend-stationary time series of turnover (SVLM) by

---

73 See Wiggins (1992), Alizadeh et al. (2002) and Chen and Daigler (2004) for detials.

74 Lobato and Velasco (2000) point out that the determination of a detrending mechanism that would allow

incorporating the procedure used by Campbell et al. (1993) that uses a 100-day backward moving average

$$SVLM = \frac{\mathrm{VO}_t}{\frac{1}{100}\sum_{i=1}^{100}\mathrm{VO}_{t-i}},$$

where VO denotes volume. This metric produces a time series that captures the change in the long run movement in trading volume (see, Brooks, 1998; Fung and Patterson, 1999). The moving average procedure is deemed to provide a reasonable compromise between computational ease and effectiveness.

5.4 Estimation Procedures

5.4.1 Estimation Methodology

Tsay and Chung (2000) have shown that regressions involving FI regressors can lead to spurious results. In particular, analyzing the bivariate regression of z_t on a constant and x_t where $z_t \sim I(d_z)$, that is integrated of order d_z, and $x_t \sim I(d_x)$ they show that the corresponding t-statistic will be divergent provided that $d_z + d_x > 0.5$ even if the two series are independent.

Moreover, in the presence of conditional heteroskedasticity, Vilasuso (2001) investigates the reliability of causality tests based on least squares. He demonstrates that when conditional heteroskedasticity is ignored, least squares causality tests exhibit considerable size distortion if the conditional variances are correlated. In addition, the inference based on a heteroskedasticity and autocorrelation consistent covariance matrix constructed under the least squares framework offers only slight improvement. Therefore, he suggests that causality tests should be carried out in the context of an empirical specification that models both the conditional means and conditional variances.

for inference on the long-memory parameter of stock volume is still an unresolved problem. Therefore, they examine consistent estimation of the long-memory parameter of volume in the frequency domain by tapering the data instead of detrending them. However, Bollerslev and Jubinski (1999) find that neither the detrending method nor the actual process of detrending affected any of their qualitative findings.

Furthermore, in many applications the sum of the estimated GARCH(1,1) parameters is often close to one, which implies integrated GARCH (IGARCH) behavior. For example, Chen and Daigler (2004) emphasize that in most cases both variables possess substantial persistence in their conditional variances. In particular, the sum of the GARCH parameters was at least 0.950. Most importantly, Baillie et al. (1996), using Monte Carlo simulations, show that data generated from a process exhibiting FIGARCH effects may be easily mistaken for IGARCH behavior. Therefore, we focus our attention on the topic of long-memory and persistence in terms of the second moments of the two variables. Consequently, we utilize a bivariate AR-FI-GARCH model to test for bidirectional causality between spread, volatility and volume.[75]

5.4.2 Dual Long-memory

Along these lines we discuss the dual long-memory time series model for the three variables and discuss its merits and properties.

Next let us define the two variables $\mathbf{y}_{i,t}$ and $\boldsymbol{\varepsilon}_{i,t}$, $i = vl, s, vo.$ the subscripts vl, s and vo mean that the first expression represents the volatility and the second one stands for the spread, the third one stands for the volume. Regarding the residual $\varepsilon_{i,t}$ we assume that it is conditionally normal with mean $\mathbf{0}$, variance $\mathbf{h}_{i,t}$.We also introduce a dummy D_t defined as: $D_t = 0$ before the introduction of electronic trading (that is 10th October 1997) and $D_t = 1$ otherwise[76].

The structure of the ARFI $(p_i, d_{m,i})$, mean equation is given by

[75] An excellent survey of major econometric work on long-memory processes and their applications in economics and finance is given by Baillie (1996). Baillie et al. (2002) and Conrad and Karanasos (2005a,b) applied the univariate dual long-memory process to inflation, and Karanasos et al. (2006) to interest rates. The bivariate dual long-memory model was introduced by Teyssière (1998). For applications to the inflation-growth link see Karanasos and Zeng (2006), for applications to the volume-volatility link see Karanasos and Kartsaklas (2008).

[76] We also test for structural breaks by employing the methodology in Bai and Perron (2003), break points in Oct 1997 were fund in all spread and volatility measurements.

$$(1-L)^{d_{m,i}}\Phi_i(L)[y_{i,t}-\Phi_{ij}(L)y_{j,t}-\Lambda_{ij}(L)y_{j,t}D_t-\mu_i^d D_t-\mu_i]=\varepsilon_{it}. \quad (5.29)$$

where, the scalar finite polynomials $\Phi_i(L)$, $\Phi_{ij}(L)$ and $\lambda_{ij}(L)$ are given by $\Phi_i(L) \triangleq 1-\sum_{k=1}^{p_i}\phi_{ik}L^k$, $\Phi_{ij}(L) \triangleq \sum_{r=1}^{p_{ij}}\phi_{ij,l}L^l$ and $\Lambda_{ij}(L) \triangleq \sum_{r=1}^{p_{ij}}\lambda_{ij,l}L^l$, $i,j = vl,s,vo$, $j \neq i$. μ_i and μ_i^d are constant and coefficient for constant dummy. The process is covariance stationary if $d_{m,i} < 0.5$ and the roots of $\Phi_i(L)$ lie outside the unit circle. The impact of $y_{j,t}$ on $y_{i,t}$ is captured by $\Phi_{ij}(L)$ for the period before the electronic trading was introduced, and $\Phi_{ij}(L)+ \Lambda_{ij}(L)$ correspond to the impact during the electronic trading period.

Further, to establish terminology and notation, the FIGARCH$(1,d_v,1)$ process is defined by[77]

$$(1-\beta_i L)(h_{i,t}-\varpi_i)=[(1-\beta_i L)-(1-L)^{d_{vi}}(1-\alpha_i L)]\varepsilon_{i,t}^2,\ i=vl,s,vo. \quad (5.30)$$

Note that the FIGARCH model is not covariance stationary. The question whether it is strictly stationary or not is still open at present (see Conrad and Haag, 2006). In the FIGARCH model conditions on the parameters have to be imposed to ensure the non-negativity of the conditional variances (see Conrad and Haag, 2006).[78]

---

77 Following Alizadeh et al. (2002), Brandt and Jones (2006) use the approximate result that if log returns are conditionally Gaussian with mean 0 and volatility h_t then the log range is a noisy linear proxy of log volatility.

78 Baillie and Morana (2007) introduce a new long-memory volatility process, denoted by Adaptive FIGARCH which is designed to account for both long-memory and structural change in the conditional variance process. One could provide an enrichment of the bivariate dual long-memory model by allowing the intercepts of the two means and variances to follow a slowly varying function as in Baillie and Morana (2007). This is undoubtedly a challenging yet worthwhile task.

Table 5.78. Mean equations: AR Lags

SPREADS	
SPRD	$1, 2, 3, 8$
ESPRD	$1, 2, 3, 4$
CSPRD	$1, 2, 3, 4, 5$
ISPRD	$1, 2, 8$
VOLATILITIES	
GVLT	$1, 2, 3, 4$
PVLT	1
VOLUME	
SVLM	$1, 4, 5, 7, 10$

Notes: The numbers represent the AR lags used in the mean equations of the model.

5.5 Empirical Results

Within the framework of the bivariate AR-FI-GARCH model[79] we will analyze the dynamic adjustments of both the conditional means and variances of spread, volatility and volume, as well as the implications of these dynamics for the direction of causality between each two variables. The estimates of the various formulations were implemented by OxMetrics with G@RCH package.

The best fitting specification is chosen according to the minimum value of the information criteria (not reported). In Table 5.78 below, we reported AR lags for the conditional means of spreads, volatilities and volume. For spreads, we choose ARFI(8, $d_{m,s}$) model for SPRD and ISPRD, we choose ARFI(4, $d_{m,s}$), ARFI(5, $d_{m,s}$) for ESPRD and CSPRD respectively. For the conditional mean of volatility, we choose ARFI(4, $d_{m,vl}$) for GVLT and ARFI(1, $d_{m,vl}$) for PVLT. Finally, SVLM has been estimated with ARFI(10, $d_{m,vo}$). We do not report the estimated AR coefficients for space considerations.

[79] Three tests aimed at distinguishing short and long-memory are implemented for the data. The statistical significance of the statistics indicates that the data are consistent with the long-memory hypothesis. (See Appendix Table 5.91)

5.5.1 The Spread-Volatility Link

The estimates of the fractional parameters for the models between spread and volatility are shown in Table 5.79. The $y_{s,t}$ and $y_{vl,t}$ columns report the results for the spread and volatility equations, respectively. Several findings emerge from this table. In all cases, the estimated value of $d_{m,s}$ is robust to the measures of volatility used. In other words, both bivariate models generated very similar mean fractional parameters $d_{m,s}$ (see $y_{s,t}$ columns). for SPRD the two long-memory mean parameters are 0.458 and 0.451 with GVLT and PVLT, respectively. ESPRD has similar estimated values of d_{ms} (0.498, 0.492), CSPRD has lower $d_{m,s}$ (0.388, 0.390) but, the corresponding values for ISPRD are higher: 0.526 and 0.524. On the other hand (see $y_{vl,t}$ columns), GVLT generated similar $d_{m,vl}$ in the models with SPRD, ESPRD and CSPRD (0.562, 0.559 and 0.554) but got a lower one with ISPRD (0.388). PVLT generated similar $d_{m,vl}$ in all four bivariate models (0.482, 0.468, 0.469 and 0.487). Generally speaking, we find that the apparent long-memory in all variables is quite resistant to the "mean shifts".

Moreover, the estimates of $d_{v,i}$ govern the long-run dynamics of the conditional heteroscedasticity with different measurements of volatility in the model, spreads generated very similar fractional variance parameters: (0.382, 0.377), (0.301, 0.300), (0.288,0.288) and (0.303, 0.304) for SPRD, ESPRD, CSPRD and ISPRD, respectively. Meanwhile, the estimated value of $d_{v,vl}$ is robust to the measures of spread used as well. In other words, all four bivariate FIGARCH models generated very similar fractional parameters. For GVLT all long-memory variance parameters are 0.481 in the models with SPRD, ESPRD and CSPRD, in the model with ISPRD $d_{v,vl}$ is 0.247. For PVLT long-memory variance parameters are 0.548 in the models with SPRD, ESPRD, CSPRD, and $d_{v,vl}$ become 0.526 with ISPRD.

Table 5.80 present the estimates of the ARCH and GARCH parameters of FIGARCH models. The $y_{s,t}$ and $y_{vl,t}$ columns report results for the spread and volatility equations respectively. Note that in all cases the GARCH coefficients satisfy the necessary and

Table 5.79. The Spread-Volatility Link, Long Memory in Mean and Variance

		SPRD		ESPRD		CSPRD		ISPRD	
		$y_{vl,t}$	$y_{s,t}$	$y_{vl,t}$	$y_{s,t}$	$y_{vl,t}$	$y_{s,t}$	$y_{vl,t}$	$y_{s,t}$
GVLT	$d_{m,i}$	0.562^{***} (0.057)	0.458^{***} (0.026)	0.559^{***} (0058)	0.498^{***} (0.049)	0.554^{***} (0.057)	0.388^{***} (0.032)	0.313^{***} (0.080)	0.526^{***} (0.051)
	$d_{v,i}$	0.481^{***} (0.068)	0.382^{***} (0.074)	0.481^{***} (0.068)	0.301^{***} (0.053)	0.481^{***} (0.084)	0.288^{***} (0.057)	0.247^{***} (0.079)	0.303^{***} (0.058)
PVLT	$d_{m,i}$	0.482^{***} (0.064)	0.451^{***} (0.029)	0.468^{***} (0.067)	0.492^{***} (0.051)	0.469^{***} (0.073)	0.390^{***} (0.035)	0.487^{***} (0.182)	0.524^{***} (0.051)
	$d_{v,i}$	0.548^{***} (0.080)	0.377^{***} (0.068)	0.548^{***} (0.079)	0.300^{***} (0.054)	0.548^{***} (0.082)	0.288^{***} (0.056)	0.526^{***} (0.062)	0.304^{***} (0.057)

Notes: The $y_{vl,t}$ and $y_{s,t}$ columns report results for the volatility and spread equations respectively.
$d_{m,i}$ and $d_{v,i}$ $(i = s, vl)$ are the long-memory parameters in mean and variance equations respectively.
***, **, * and $^{\circ}$ denotes significance at the 0.01, 0.05, 0.10 and 0.15 levels respectively.
The numbers in parentheses are standard errors.

Table 5.80. The Spread-Volatility Link, ARCH and GARCH Coefficients

		SPRD		ESPRD		CSPRD		ISPRD	
		$y_{vl,t}$	$y_{s,t}$	$y_{vl,t}$	$y_{s,t}$	$y_{vl,t}$	$y_{s,t}$	$y_{vl,t}$	$y_{s,t}$
GVLT	α	-0.123 (0.176)	0.344° (0.074)	-0.119 (0.173)	0.073 (0.199)	-0.122 (0.173)	–	0.879^{***} (0.051)	–
	β	0.177 (0.190)	0.604^{**} (0.235)	0.180 (.187)	0.250 (0.205)	0.181 (0.187)	0.175^{**} (0.080)	0.720^{***} (0.059)	0.184^{***} (0.071)
PVLT	α	–	0.345° (0.231)	–	0.094 (0.194)	–	–	–	–
	β	0.332^{***} (0.117)	0.596^{**} (0.237)	0.332^{***} (0.117)	0.271 (0.196)	0.333^{***} (0.120)	0.173^{**} (0.080)	-0.024 (0.091)	0.185^{***} (0.072)

Notes: The $y_{vl,t}$ and $y_{s,t}$ columns report results for the volatility and spread equations respectively.
α and β are the ARCH and GARCH parameters.
***, **, * and $^{\circ}$ denotes significance at the 0.01, 0.05, 0.10 and 0.15 levels respectively.
The numbers in parentheses are standard errors.

sufficient conditions for the non-negativitiy of the conditional variances (see Conrad and Haag, 2006).

We employ the bivariate AR-FI-GARCH model with lagged values of one variable included in the mean equation of the other variable to test for bidirectional causality. The estimated coefficients $\phi_{ij,l}$ and $\lambda_{ij,l}$ are defined in Equation (5.29), which capture the possible feedback between the two variables. Recall that the relationship between the two variables before the adoption of electronic trading system is captured by $\phi_{ij,l}$, while $(\phi_{ij,l} + \lambda_{ij,l})$ captures the link commencing with the introduction of electronic trading.

Table 5.81. The Spread-Volatility Link, Cross Effects

		SPRD		ESPRD		CSPRD		ISPRD	
		$y_{vl,t}$	$y_{s,t}$	$y_{vl,t}$	$y_{s,t}$	$y_{vl,t}$	$y_{s,t}$	$y_{vl,t}$	$y_{s,t}$
GVLT	$\phi_{ij,l}$	-0.922^{**} (0.470) $l=3$	0.017^{***} (0.006) $l=2$	-0.329^{**} (0.134) $l=3$	0.073^{***} (0.022) $l=2$	-0.554^{**} (0.243) $l=3$	0.021^{*} (0.012) $l=2$	4.029^{***} (0.240) $l=0$	0.032^{***} (0.010) $l=2$
	$\lambda_{ij,l}$	-0.082 (0.630) $l=3$	-0.002 (0.006) $l=2$	-0.162 (0.340) $l=3$	-0.030° (0.019) $l=2$	-0.265 (0.615) $l=3$	-0.006 (0.010) $l=2$	1.735^{***} (0.557) $l=0$	-0.013° (0.008) $l=2$
PVLT	$\phi_{ij,l}$	-0.417^{**} (0.176) $l=2$	$0.039^{**\dagger}$ (0.016) $l=2$	-0.132^{***} (0.048) $l=2$	0.193^{**} (0.062) $l=2$	-0.236^{***} (0.091) $l=2$	0.047° (0.033) $l=2$	1.668^{***} (0.099) $l=0$	0.084^{***} (0.027) $l=2$
	$\lambda_{ij,l}$	0.096 (0.292) $l=2$	-0.017 (0.016) $l=2$	-0.033 (0.142) $l=2$	-0.098^{*} (0.050) $l=2$	-0.093 (0.239) $l=2$	-0.022 (0.028) $l=2$	1.394^{***} (0.191) $l=0$	-0.044^{*} (0.023) $l=2$

Notes: The $y_{vl,t}$ and $y_{s,t}$ columns report results for the volatility and spread equations respectively.
$\phi_{ij,l}$ and $\lambda_{ij,l}$ are the parameter estimates of the cross effects and coefficients for slope dummy.
***, **, * and $^{\circ}$ denotes significance at the 0.01, 0.05, 0.10 and 0.15 levels respectively.
$\dagger$ lag 1 of pvlt (insignificant) also included.
The numbers in parentheses are standard errors.

Table 5.81 reports bidirectional feedback between spread and volatility. The $y_{s,t}$ and $y_{vl,t}$ columns report results for the spread and volatility equations respectively. $\phi_{ij,l}$ are significant in all models, however, we only get significant $\lambda_{ij,l}$ when ESPRD is the dependent variable or in the models between ISPRD and volatilities. Both GVLT and PVLT have positive effect on all spread measures, in particular, this positive effect become stronger after electronic trading was adopted in 1997 on both ESPRD and ISPRD (see $y_{s,t}$ columns). On another hand (see $y_{vl,t}$ columns), spreads include SPRD, ESPRD and CSPRD have negative effect on volatilities. On contrary, ISPRD has positive effect on both GVLT and PVLT and this positive effect become higher in magnitude during electronic trading period. Table 5.82 below gives an overview of the spread-volatility link over the two different periods.

5.5.2 The Spread-Volume Link

The estimates of the fractional parameters for the models between spread and volume are shown in Table 5.83. The $y_{s,t}$ and $y_{vo,t}$ columns report results for the spread and volume equations respectively. Several findings emerge from this table. The estimated value of $d_{m,s}$ are equal to 0.491, 0.545, 0.424 and 0.347 for SPRD ESPRD CSPRD and ISPRD

Table 5.82. The Spread-Volatility Link, Overview

Sample:	SPRD	ESPRD	CSPRD	ISPRD
Panel A. The effect of Volatility on Spread				
GVLT	Positive	Positive^{la}	Positive	Positive^{la}
PVLT	Positive	Positive^{la}	Positive	Positive^{la}
Panel B. The impact of Spread on Volatility				
GVLT	Negative	Negative	Negative	Positive^{ha}
PVLT	Negative	Negative	Negative	Positive^{ha}

Notes: ^{la} The positive effect is less in magnitude after electronic trading was adopted in year 1997.
^{ha} The positive effect is higher in magnitude after electronic trading was adopted in year 1997.

Table 5.83. The Spread-Volume Link, Long Memory in Mean and Variance

	SPRD		ESPRD		CSPRD		ISPRD	
	$y_{vo,t}$	$y_{s,t}$	$y_{vo,t}$	$y_{s,t}$	$y_{vo,t}$	$y_{s,t}$	$y_{vo,t}$	$y_{s,t}$
$d_{m,i}$	0.179^{***}	0.491^{***}	0.178^{***}	0.545^{***}	0.178^{***}	0.424^{***}	0.179^{***}	0.347^{***}
	(0.042)	(0.027)	(0.042)	(0.040)	(0.042)	(0.032)	(0.042)	(0.025)
$d_{v,i}$	0.168	0.388^{***}	0.168	0.303^{***}	0.166	0.279^{***}	0.170	0.305^{***}
	(0.130)	(0.070)	(0.129)	(0.053)	(0.127)	(0.051)	(0.132)	(0.059)

Notes: The $y_{vo,t}$ and $y_{s,t}$ columns report results for the volume and spread equations respectively.
$d_{m,i}$ and $d_{v,i}$ $(i = s, vo)$ are the long-memory parameters in mean and variance equations respectively.
***, **, * and $^{\circ}$ denotes significance at the 0.01, 0.05, 0.10 and 0.15 levels respectively.
The numbers in parentheses are standard errors.

respectively. On the other hand, SVLM generated similar $d_{m,vo}$ in the models with different measurement of spread (0.179, 0.178, 0.178 and 0.179 for SPRD, ESPRD,CSPRD and ISPRD respectively). Generally speaking, we find that the apparent long-memory in all variables is quite resistant to the “mean shifts”.

Moreover, The estimates of $d_{v,i}$’s govern the long-run dynamics of the conditional heteroscedasticity. The estimated value of $d_{v,s}$ are equal to 0.388, 0.303, 0.279 and 0.305 for SPRD ESPRD CSPRD and ISPRD respectively (see $y_{s,t}$ columns). In all cases the estimated value of $d_{v,vo}$ is robust to the measures of spreads used (see $y_{vo,t}$ columns). In other words, FIGARCH models generated very similar $d_{v,vo}$ ’s fractional parameters, SVLM’s long-memory variance parameters are 0.168 in models with SPRD, ESPRD, 0.166 with CSPRD and 0.170 with ISPRD.

Table 5.84. The Spread-Volume Link, ARCH and GARCH Coefficients

	SPRD		ESPRD		CSPRD		ISPRD	
	$y_{vo,t}$	$y_{s,t}$	$y_{vo,t}$	$y_{s,t}$	$y_{vo,t}$	$y_{s,t}$	$y_{vo,t}$	$y_{s,t}$
α	–	0.389^{*} (0.214)	–	–	–	–	–	–
β	0.388^{*} (0.153)	0.633^{***} (0.219)	0.388^{*} (0.152)	0.303^{*} (0.187)	0.385^{**} (0.151)	0.143^{**} (0.073)	0.391^{**} (0.154)	0.177^{**} (0.074)

Notes: The $y_{vo,t}$ and $y_{s,t}$ columns report results for the volume and spread equations respectively.
α and β are the ARCH and GARCH parameters.
***, **, * and $^{\circ}$ denotes significance at the 0.01, 0.05, 0.10 and 0.15 levels respectively.
The numbers in parentheses are standard errors.

Table 5.84 present the estimates of the ARCH and GARCH parameters of FIGARCH models. The $y_{s,t}$ and $y_{vo,t}$ columns report results for the spread and volume equations respectively. Note that in all cases, the GARCH coefficients satisfy the necessary and sufficient conditions for the non-negativitiy of the conditional variances (see Conrad and Haag, 2006).

We employ the bivariate AR-FI-GARCH model with lagged values of one variable included in the mean equation of the other variable to test for bidirectional causality. The estimated coefficients $\phi_{ij,l}$ and $\lambda_{ij,l}$ are defined in Equation (5.29), which capture the possible feedback between the two variables. Recall that the relationship between the two variables before the adoption of electronic trading system is captured by $\phi_{ij,l}$, while $\left(\phi_{ij,l} + \lambda_{ij,l}\right)$ captures the link commencing with the introduction of electronic trading.

Table 5.85 reports bidirectional feedback between spread and volatility. The $y_{s,t}$ and $y_{vo,t}$ columns report results for the spread and volatility equations respectively. $\phi_{ij,l}$ are significant in all models, and $\lambda_{ij,l}$ are significant in all spread equations ($y_{s,t}$ columns). Volume has positive effect on spread. Interestingly, in all cases, the information criteria chooses "0" as the order of lags for SVLM, this result indicates that volume and spreads tend to move toward the same direction simultaneously, We also noticed that this effect become higher in Magnitude after electronic trading was adopted in year 1997. On the other hand, spreads have negative influence on SVLM. Table 5.86 below gives an overview of the spread-volume link over the two different periods considered.

Table 5.85. The Spread-Volume Link, Cross Effects

	SPRD		ESPRD		CSPRD		ISPRD	
	$y_{vo,t}$	$y_{s,t}$	$y_{vo,t}$	$y_{s,t}$	$y_{vo,t}$	$y_{s,t}$	$y_{vo,t}$	$y_{s,t}$
$\phi_{ij,l}$	-0.075^{*}	0.149^{***}	-0.027^{*}	0.457^{***}	-0.049^{*}	0.231^{***}	-0.059^{*}	0.188^{***}
	(0.042)	(0.015)	(0.055)	(0.049)	(0.028)	(0.024)	(0.034)	(0.024)
	$l=6$	$l=0$	$l=6$	$l=0$	$l=6$	$l=0$	$l=6$	$l=0$
$\lambda_{ij,l}$	0.060	0.234^{***}	0.020	0.277^{***}	0.034	0.172^{***}	0.045	0.167^{***}
	(0.043)	(0.045)	(0.016)	(0.107)	(0.030)	(0.057)	(0.035)	(0.053)
	$l=6$	$l=0$	$l=6$	$l=0$	$l=6$	$l=0$	$l=6$	$l=0$

Notes: The $y_{vo,t}$ and $y_{s,t}$ columns report results for the volume and spread equations respectively.
$\phi_{ij,l}$ and λ are the parameter estimates of the cross effects and coefficients for slope dummy.
***, **, * and $^{\circ}$ denotes significance at the 0.01, 0.05, 0.10 and 0.15 levels respectively.
The numbers in parentheses are standard errors.

Table 5.86. The Spread-Volume Link, Overview

Sample:	SPRD	ESPRD	CSPRD	ISPRD
Panel A. The effect of Volume on Spread				
SVLM	Positive^{ha}	Positive^{ha}	Positive^{ha}	Positive^{ha}
Panel B. The impact of Spread on Volume				
SVLM	Negative	Negative	Negative	Negative

Notes: ^{ha} Effect is higher in magnitude after electronic trading was adopted in year 1997.

5.5.3 The Volume-Volatility Link

The estimates of the fractional parameters for the models between spread and volume are shown in Table 5.87. The $y_{vo,t}$ and $y_{vl,t}$ columns report results for the volume and volatility equations respectively. Several findings emerge from this table. The estimated value of $d_{m,vo}$ is robust to the measures of volatility used (equal to 0.150 with both GVLT and PVLT). On the other hand, $d_{m,vl}$ are equal to 0.474 and 0.333 for models of GVLT and PVLT respectively.

Moreover, The estimates of $d_{v,i}$'s govern the long-run dynamics of the conditional heteroscedasticity. The estimated value of $d_{v,vo}$ are equal to 0.577 and 0.578 for GVLT and PVLT respectively (see $y_{vo,t}$ columns), which imply that the estimated value of $d_{v,vo}$ is robust to the measures of volatility used. long-memory variance parameters $d_{v.vl}$ for GVLT and PVLT are 0.493 and 0.540 respectively.

Table 5.87. The Volume-Volatility Link, Long Memory in Mean and Variance

	GVLT		PVLT	
	$y_{vo,t}$	$y_{vl,t}$	$y_{vo,t}$	$y_{vl,t}$
$d_{m,i}$	0.150^{***} (0.046)	0.474^{***} (0.051)	0.150^{***} (0.046)	0.333^{***} (0.147)
$d_{v,i}$	0.577^{***} (0.136)	0.493^{***} (0.059)	0.578^{***} (0.135)	0.540^{***} (0.075)

Notes: The $y_{vo,t}$ and $y_{vl,t}$ columns report results for the volume and volatility equations respectively.
$d_{m,i}$ and $d_{v,i}$ $(i = vo, vl)$ are the long-memory parameters in mean and variance equations respectively.
***, **, * and $^{\circ}$ denotes significance at the 0.01, 0.05, 0.10 and 0.15 levels respectively.
The numbers in parentheses are standard errors.

Table 5.88. The Volume-Volatility Link, ARCH and GARCH Coefficients

	GVLT		PVLT	
	$y_{vo,t}$	$y_{vl,t}$	$y_{vo,t}$	$y_{vl,t}$
α	–	0.165 (0.157)	–	–
β	0.722^{***} (0.096)	0.204 (0.162)	0.723^{***} (0.096)	0.327^{***} (0.121)

Notes: The $y_{vo,t}$ and $y_{vl,t}$ columns report results for the volume and volatility mean equations respectively.
α and β are the ARCH and GARCH parameters.
***, **, * and $^{\circ}$ denotes significance at the 0.01, 0.05, 0.10 and 0.15 levels respectively.
The numbers in parentheses are standard errors.

Table 5.88 present the estimates of the ARCH and GARCH parameters of FIGARCH models. The $y_{vo,t}$ and $y_{vl,t}$ columns report results for the volume and volatility equations respectively. Note that in all cases, the GARCH coefficients satisfy the necessary and sufficient conditions for the non-negativitiy of the conditional variances (see Conrad and Haag, 2006).

We employ the bivariate AR-FI-GARCH model with lagged values of one variable included in the mean equation of the other variable to test for bidirectional causality. The estimated coefficients $\phi_{ij,l}$ and $\lambda_{ij,l}$ are defined in Equation (5.29), which capture the possible feedback between the two variables. Recall that the relationship between the two variables before the adoption of electronic trading system is captured by $\phi_{ij,l}$, while $\left(\phi_{ij,l} + \lambda_{ij,l}\right)$ captures the link commencing with the introduction of electronic trading.

Table 5.89 reports bidirectional feedback between volume and volatility. The $y_{vo,t}$ and $y_{vl,t}$ columns report results for the volume and volatility equations respectively. SVLM shows no influence on GVLT, however, it exhibited positive impact on PVLT($\phi_{ij,l} = 0.082$). On the other hand, $\phi_{ij,l}$ are positive and significant and $\lambda_{ij,l}$ are negative and significant in the $y_{vo,t}$ columns, which indicates positive impacts from both GVLT and PVLT on SVLM and these positive effects is less in magnitude during electronic trading period. Table 5.90 below gives an overview of the volume-volatility link over the two different periods.

Table 5.89. The Volume-Volatility Link, Cross Effects

	GVLT		PVLT	
	$y_{vo,t}$	$y_{vl,t}$	$y_{vo,t}$	$y_{vl,t}$
$\phi_{ij,l}$	0.014^{*} (0.019) $l=2$	0.044 (0.112) $l=2$	0.035^{*} (0.019) $l=2$	0.082^{**} (0.042) $l=5$
$\lambda_{ij,l}$	-0.013^{*} (0.019) $l=2$	-1.152 (1.185) $l=2$	-0.035^{*} (0.019) $l=2$	-0.174 (0.340) $l=5$

Notes: The $y_{vo,t}$ and $y_{vl,t}$ columns report results for the volume and volatility equations respectively.
$\phi_{ij,l}$ and $\lambda_{ij,l}$ are the parameter estimates of the cross effects and coefficients for slop dummy.
***, **, * and $^{\circ}$ denotes significance at the 0.01, 0.05, 0.10 and 0.15 levels respectively.
The numbers in parentheses are standard errors.

Table 5.90. The Volume-Volatility Link, Overview

Sample:	GVLT	PVLT
Panel A. The effect of Volume on Volatility		
SVLM	-	Positive
Panel B. The impact of Volatility on Volume		
SVLM	Positiv^{la}	Positive^{la}

Notes: ^{la} The positive effect is less in magnitude after electronic trading was adopted in year 1997.

5.6 Conclusion

In this study we investigate two-way causality relationships between spread, volatility and volume in FTSE100 stock index over the period 1992–2004 by using bivariate AR-FI-GARCH model and multiple measurements of risk and spread. Measurements of spread include relative bid-ask spread, effective bid-ask spread, the inventory cost component of the bid-ask spread and the information cost component of the bid-ask spread. Risk is measured by two measurements of price volatility: the close-to-close volatility and the range-based volatility. we also take the impact of electronic trading into account.

Our results are in line with the general agreement that volatility positively impacts bid-ask spread, more specifically, we find evidence supporting Roll (1984) and Karpoff (1986) that market maker widen bid-ask spread in order to compensate the cost of inventory, since our results show that inventory cost component of the bid-ask spread is positively affected by volatility. We also find that the information cost component of the bid-ask spread gets higher when risk of stock index increases, this finding supports the claim of Admati and Pfleiderer (1988) and Kim and Verrecchia (1994) that higher risk associates with more informed trading and market makers require premium for their disadvantage in the information asymmetry. Furthermore, we find that the impact of volatility on effective spread and informational cost is weakened by the adoption of electronic trading system. We also find evidence supporting positive correlation between volume and volatility. Interestingly, we also observed co-movement of volume and spread, this may be due to their same reaction on the impact of volatility(both positively affected by volatility with order of lags equal to 2) and the impact from volume on spread is too weak to compare with that of volatility (see Appendix Table 5.92 and 5.93).

On the other hand, both volatility and volume are negatively affected by the relative spread and the effective spread. When we classify the components of bid-ask spread in this causality investigation, we find that only inventory cost component of spread has negative effect on volatility. In contrast, information component of spread positively impacts volatility and this effect become stronger since electronic trading has been used. In addition, all measurements of spread components have negative impact on volume. These results support the argument that speculations generate volatility in the market and higher transaction cost benefit the stability of the market.

5.A Appendix

The KPSS test (Kwiatkowski et al., 1992), Lo's modified R/S test (Lo,1991) and the HML test (Harris, McCabe and Leybourne, 2008) were conducted in order to distinguishing long and short memory for spreads, volatilities and volumes, "The KPSS and R/S tests are alternative implementations based on the partial sums of the series, and have nonstandard limiting distributions. The HML test is asymptotically N(0,1), and is based on the long-range autocovariances." (Davidson, 2009, p.50). As seen in Table 5.91 , all data series include spreads, volatilities and volume are exhibit long memory property.

Table 5.91. Long Memory Tests

	KPSS test	Lo's RS test	HML test†
SPREADS			
SRD	92.808^{***}	18.901^{***}	7.641^{***}
ESPRD	58.436^{***}	14.605^{***}	5.759^{***}
ISPRD	57.474^{***}	14.492^{***}	5.695^{***}
CSPRD	58.653^{***}	14.625^{***}	5.808^{***}
VOLATILITIES			
GVLT	25.313^{***}	9.474^{***}	3.904^{***}
PVLT	25.857^{***}	9.681^{***}	3.952^{***}
VOLUME			
SVLM	0.675^{**}	2.828^{***}	-0.485

Notes: The table reports values of long memory tests statistic. †truncation parameters $C = 1$ and $L = 0.66$. ***, **and * denotes significance at the 0.01, 0.05 and 0.10 levels respectively.

Table 5.92 and 5.93 reports the ARFI-FIGARCH estimation of spread on volatility and volume[80]. In all cases, $\phi_{vo,l}$ are insignificant.

[80] The models on ESPRD have no convergence.

Table 5.92. ARFI-FIGARCH Estimation: Spreads on GVLT and SVLM

	SPRD	CSPRD	ISPRD
	Panel A: Mean Equations		
$\phi_{vl,l}$	0.018^{***}	0.024^{**}	0.006
	(0.006)	(0.012)	(0.010)
	$l=2$	$l=2$	$l=2$
$\lambda_{vl.l}$	-0.004	-0.009	-0.003
	(0.006)	(0.011)	(0.010)
$\phi_{vo,l}$	-0.008	-0.006	-0.015
	(0.020)	(0.024)	(0.018)
	$l=3$	$l=2$	$l=3$
$\lambda_{vo.l}$	-0.027	0.058	-0.029
	(0.042)	(0.053)	(0.042)
d_m	0.450^{***}	0.385^{***}	0.337^{***}
	(0.025)	(0.032)	(0.037)
	Panel B: Variance Equations		
α	0.343°	–	–
	(0.219)		
β	0.619^{***}	0.152^{*}	0.221^{**}
	(0.231)	(0.082)	(0.099)
d_v	0.400^{***}	0.282^{***}	0.339^{***}
	(0.080)	(0.058)	(0.088)

Notes: $\phi_{i,l}$ and $\lambda_{i,l}$ are the parameter estimates of the cross effects and coefficients for slop dummy.
d_m and d_v are the long-memory parameters in mean and variance equations respectively.
α and β are the ARCH and GARCH parameters.
***, **, * and $^{\circ}$ denotes significance at the 0.01, 0.05, 0.10 and 0.15 levels respectively.
The model on ESPRD has no convergence
The numbers in parentheses are standard errors.

Table 5.93. ARFI-FIGARCH Estimation: Spreads on PVLT and SVLM

	SPRD	CSPRD	ISPRD
	Panel A: Mean Equations		
$\phi_{vl,l}$	0.044^{***} (0.017) $l=2$	0.058^{*} (0.035) $l=2$	0.013 (0.028) $l=2$
$\lambda_{vl.l}$	-0.012° (0.008)	-0.034 (0.030)	-0.018 (0.028)
$\phi_{vo,l}$	-0.003 (0.016) $l=2$	-0.009 (0.024) $l=2$	-0.021 (0.027) $l=3$
$\lambda_{vo.l}$	0.056 (0.045)	0.064 (0.054)	-0.019 (0.045)
d_m	0.455^{***} (0.037)	0.387^{***} (0.034)	0.356^{***} (0.039)
	Panel B: Variance Equations		
α	0.364^{*} (0.224)	–	–
β	0.625^{***} (0.227)	0.150^{*} (0.081)	0.182^{**} (0.085)
d_v	0.392^{***} (0.074)	0.281^{***} (0.056)	0.308^{***} (0.067)

Notes: $\phi_{ij,l}$ and $\lambda_{i,l}$ are the parameter estimates of the cross effects and coefficients for slop dummy.
d_m and d_v are the long-memory parameters in mean and variance equations respectively.
α and β are the ARCH and GARCH parameters.
***, **, * and $^{\circ}$ denotes significance at the 0.01, 0.05, 0.10 and 0.15 levels respectively.
The model on ESPRD has no convergence
The numbers in parentheses are standard errors.

REFERENCES

Acemoglu, D. and J. Robinson, 2006, *Economic Origins of Dictatorship and Democracy*. Cambridge University Press, Boston.

Acemogly, D., 2009, *Introduction to Modern Economic Growth*, Princeton University Press.

Admati, A. R. and P. Pfleiderer, 1988, A theory of intraday patterns: volume and price variability. *Review of Financial Studies* 1, 3–40.

Aiolfi, M., L. Catão and A. Timmermann, 2010, Common Factors in Latin America's Business Cycles. *Journal of Development Economics*, Forthcoming**.**

Alesina, A. and L. H. Summers, 1993, Central Bank Independence and Macroeconomic Performance: Some Comparative Evidence, *Journal of Money, Credit and Banking* 25(2), 151-162.

Alizadeh, S., M. Brandt, and F. Diebold, 2002, Range-based Estimation of Stochastic Volatility Models. *Journal of Finance* 57, 1047-1091.

Alston, L. and A. A. Gallo, 2007, Electoral Fraud, the Rise of Peron and Demise of Checks and Balances in Argentina. Mimeo, University of Colorado.

Asteriou, D. and S. Price, 2001, Political Instability and Economic Growth: UK Time Series Evidence. *Scottish Journal of Political Economy* 48, 383-399.

Atkins, A.B. and E.A. Dyl, 1997, Transactions Costs and Holding Periods for Common Stocks. *Journal of Finance* Vol. 52, pp. 309–25.

Bai, J. and P. Perron, 2003, Computation and Analysis of Multiple Structural Change Models. *Journal of Applied Econometrics* 18, 1-22.

Baillie, R., Bollerslev, T., and H. Mikkelsen, 1996, Fractionally Integrated Generalized Autoregressive Conditional Heteroskedasticity. *Journal of Econometrics* 74, 3–30.

Baillie, R. T., C. Chung and M. Tieslau, 1996, Analyzing Inflation by the Fractionally Integrated ARFIMA-GARCH Model. *Journal of Applied Econometrics* 11, 23-40.

Baillie, R. T., Y. W. Han and T. Kwon, 2002, Further Long Memory Properties of Inflation Shocks. *Southern Economic Journal* 68, 496-510.

Ball, L. N., G. Mankiw, and D. Romer, 1988, The new Keynesian Economics and the Output-inflation Trade-off. *Brookings Papers on Economic Activity* 1: 165.

Ball, L., 1992, Why Does High Inflation Raise Inflation Uncertainty? *Journal of Monetary Economics* 29, 371-388.

Banks, A., 2005, Cross-National Time-Series Data Archive. Jerusalem: Databanks International. http://www.databanks.sitehosting.net.

Beck, T., R. Levine and N. Loayza, 2000, Finance and Sources of Growth. *Journal of Financial Economics* 58, 261–300.

Bekaert, G., C. R. Harvey and C. Lundblad, 2006, Growth Volatility and Equity Market Liberalization. *Journal of International Money and Finance* 25, 370-403.

Benston, G. and R. Hagerman, 1974, Determinants of Bid–asked Spreads in the Over-the-counter Market. *Journal of Financial Economics*, pp.351–364.

Bernanke, B., 1983, Irreversibility, Uncertainty, and Cyclical Investment. *Quarterly Journal of Economics* 98, no. 185106.

Bhar, R. and S. Hamori, 2003, Alternative Characterisation of the Volatility in the Growth Rate of Real GDP. *Japan and the World Economy* 15, no. 2: 22331.

Black, F., 1987, *Business Cycles and Equilibrium*. New York: Basil Blackwell.

Blackburn, K., 1999, Can Stabilisation Policy Reduce Long-run Growth? *Economic Journal* 109, no. 1: 6777.

Blackburn, K., and R. Galinder, 2003, Growth, volatility and Learning. *Economics Letters* 79, no. 3:1721.

Blackburn, K., and A. Pelloni, 2004, On the Relationship Between Growth and Volatility. *Economics Letters* 83, no. 1: 1237.

Blackburn, K., and A. Pelloni, 2005, Growth, Cycles and Stabilisation Policy. *Oxford Economic Papers* 57, no. 2: 26282.

Bollerslev, T., and J. M. Wooldridge, 1992, Quasi-maximum Likelihood Estimation and Inference in Dynamic Models with Time Varying Covariances. *Econometric Reviews* 11, no. 2: 14372.

Bollerslev, T., and D. Jubinski, 1999, Equity Trading Volume and Volatility: Latent Information Arrivals and Common Long-run Dependencies. *Journal of Business & Economic Statistics*, 1, 9-21.

Bollerslev, T. and M. T. Melvin, 1994, Bid–ask Spreads and the Volatility in the Foreign Exchange Market. *Journal of International Economics* 36, 355–372.

Booth, G. and U. Gurun, 2008, Volatility Clustering and the Bid–ask Spread: Exchange Rate Behaviour in Early Renaissance Florence. *Journal of Empirical Finance* 15 (2008) 131–144.

Bordo, M., B. Eichengreen, D. Klingebiel, and M. S. Martinez-Peria, 2001, Is the Crisis Problem Growing More Severe? *Economic Policy* 16(32), 51 - 82.

Brooks, C., 1998, Predicting Stock Index Volatility: can Market Volume Help? *Journal of Forecasting* 17, 59-80.

Brooks, R. D., R. W. Faff and M. D. McKenzie, 2000, A Multi-Country Study of Power ARCH Models and National Stock Market Returns, *Journal of International Money and Finance* 19, 377-397.

Briault, C., 1995, The Costs of Inflation. *Bank of England Quarterly Bulletin* February: 3345.

Brunner, A. D. and G. D. Hess, 1993, Are Higher Levels of Inflation Less Predictable? A State-Dependent Conditional Heteroscedasticity Approach. *Journal of Business & Economic Statistics* 11, 187-197.

Bryant, H. and M. Haigh, 2004, Bid-ask Spreads in Commodity Futures Markets. *Applied Financial Economics* 14:923-936.

Campbell, J., Grossman, S., and Wang, J., 1993, Trading volume and serial correlation in stock returns. *Quarterly Journal of Economics* 108, 905-939.

Campos, N. and J. Nugent, 2002, Who is Afraid of Political Instability? *Journal of Development Economics* 67, 157-172.

Campos, N. and J. Nugent, 2003, Aggregate Investment and Political Instability: An Econometric Investigation. *Economica* 70(3), 533-549.

Campos, N. and M. Karanasos, 2008, Growth, Volatility and Instability: Non-Linear Time-Series Evidence for Argentina. *Economics Letters* 100(1), 135-137.

Campos, N., M. Karanasos and B. Tan, 2008, From Riches to Rags, and Back? Explaining the Growth Trajectory of Argentina since the 1890s, *CEPR DP* 6897.

Campos, N., M. Karanasos and M. Karoglou, 2009, Apocalypse When? An Investigation of Structural Break in the Economic History of Argentina, mimeo.

Caporale, T., and B. McKiernan, 1996, The Relationship Between Output Variability and Growth: Evidence from Post War UK Data. *Scottish Journal of Political Economy* 43, no. 2: 22936.

Caporale, T., and B. McKiernan, 1998, The Fischer Black Hypothesis: Some Time-series Evidence. *Southern Economic Journal* 64, no. 3: 76571.

Chen, Z., and R Daigler, 2004, Linear and Nonlinear Interaction among Futures Volatility, Institutional Traders, and the General Public. Unpublished paper, Department of Finance, Florida International University.

Chen, S.W., C. H. Shen and Z. Xie, 2006, Nonlinear Relationship between Inflation and Inflation Uncertainty in Taiwan, *Applied Economics Letters*, 13: 8, 529-533.

Chen, S.W., C. H. Shen and Z. Xie, 2008, Evidence of a Nonlinear Relationship between Inflation and Inflation Uncertainty: The Case of the Four Little Dragons, *Journal of Policy Modeling* 30, 363–376.

Clarida, R., J. Gal and M. Gertler, 1999, The Science of Monetary Policy: A New Keynesian Perspective, *Journal of Economic Literature* 37(4), 1661-1707.

Conrad, C., and B. Haag, 2006, Inequality Constraints in the FIGARCH model. *Journal of Financial Econometrics* 3, 413-449.

Conrad C., F. Jiang and M. Karanasos, 2006, Modelling and Predicting Exchange Rate Volatility Via Power ARCH Models: The Role of Long-Memory, Unpublished paper, University of Mannheim.

Conrad, C. and M. Karanasos, 2005a, On the Inflation-Uncertainty Hypothesis in the USA, Japan and the UK: A Dual Long Memory Approach. *Japan and the World Economy* 17, 327-343.

Conrad, C. and M. Karanasos, 2005b, Dual Long Memory in Inflation Dynamics Across Countries of the Euro Area and the Link Between Inflation Uncertainty and Macroeconomic Performance, *Studies in Nonlinear Dynamics and Econometrics* 9, Article 5.

Conrad, C., M. Karanasos, 2010. Modelling the Link between US Inflation and Output: the Importance of the Uncertainty Channel. University of Heidelberg Department of Economics Discussion Paper No. 507

Conrad, C., M. Karanasos and N. Zeng, 2007, Fractionally Integrated APARCH Modelling of Stock Market Volatility: A Multi Country Study, Unpublished Paper, University of Mannheim.

Conrad, C., M. Karanasos, and N. Zeng, 2010, The Link between Macroeconomic Performance and Variability in the UK. *Economics Letters* 106, 154–157.

Constantinides, G., 1986, Capital Market Equilibrium with Transactions Costs, *Journal of Political Economy* Vol. 94, pp. 842–62.

Copeland, T. and D. Galai, 1983, Information Effects on the Bid-ask Spread, *Journal of Finance*, Vol. 28, 1457-1469

Cortes C. R., 2009, *The Political Economy of Argentina in the Twentieth Century*. Cambridge University Press, Cambridge.

Crouzille, C.and L. Lepetit and A. Tarazi, 2004, Bank Stock Volatility, News and Asymmetric Information in Banking: an Empirical Investigation. *Journal of Multinational Financial Management* Vol 14, Issues 4-5, Pages 443-461

Cukierman, A. and A. Meltzer, 1986, A Theory of Ambiguity, Credibility, and Inflation Under Discretion and Asymmetric Information. *Econometrica* 54, 1099-1128.

Dotsey, M. and P. Sarte, 2000, Inflation Uncertainty and Growth in a Cash-in-advance Economy. *Journal of Monetary Economics* 45(3), pp. 631-655.

Diaz-Alejandro, C., 1985, Argentina, Australia and Brazil Before 1929. In Platt, D. C. M. and di Tella, Guido (Eds.), *Argentina, Australia and Canada. Studies in Comparative Development 1870-1965*. St. Martin's Press, New York.

Ding, D and C. Charoenwong, 2003, Bid-ask Spreads, Volatility, Quote Revisions, and Trades of the Thinly Traded Futures Contracts. *The Journal of Futures Markets* 23, No. 5, 455–486

Ding, D. and B. Chong, 1997, Simex Nikkei Futures Spreads and their Determinants. *Advances in Pacific Basin Financial Markets* 3:39-53.

Ding, Z., C.W.J. Granger and R. Engle, 1993, A Long Memory Property of Stock Market Returns and a New Model. *Journal of Empirical Finance* 1, 83-106.

Dooley, M. P., 1996, Tobin Tax: Good Theory, Weak Evidence, Questionable Policy, in ul Haq et al. (eds.), *The Tobin tax: Coping with financial volatility*, New York, Oxford: Oxford University Press, 83-106.

Dueker, M. and G. Kim, 1999, A Monetary Policy Feedback Rule in Korea's Fast-growing Economy, *Journal of International Financial Markets, Institutions and Money* 9, 19–31.

Durlauf, S., P. Johnson and J. Temple, 2005, Growth Econometrics. In Aghion, P. and S. Durlauf (Eds.), *Handbook of Economic Growth*, North-Holland.

Easley, D., M. O'Hara, 1992, Time and the Process of Security Price Adjustment. *Journal of Finance* 47, 577–605.

Fecht, F., K. Huang and A. Martin, 2008, Financial Intermediaries, Markets, and Growth. *Journal of Money, Credit and Banking* 40(4), 701-720.

Fornari, F. and A. Mele, 1997, Weak Convergence and Distributional Assumptions for a General Class of Nonlinear ARCH Models. *Econometric Reviews* 16, 205-229.

Fornari, F. and A. Mele, 2001, Recovering the Probability Density Function of Asset Prices Using GARCH as Diffusion Approximations. *Journal of Empirical Finance* 8, 83-110.

Foster, F. G. and S. Viswanathan, 1993, Variation in Trading Volume, Return Volatility, and Trading Costs: Evidence on Recent Price Formation Models. *Journal of Finance* Vol. 48, 187-211.

Fountas, S., A. Ioannidi and M. Karanasos, 2004, Inflation, Inflation Uncertainty, and a Common European Monetary Policy. *Manchester School* 2, 221-242.

Fountas, S. and M. Karanasos, 2007, Inflation, Output Growth, and Nominal and Real Uncertainty: Empirical Evidence for the G7. *Journal of International Money and Finance* 26, 229-250.

Fountas, S. and M. Karanasos, 2008, Are Economic Growth and the Variability of the Business Cycle Related? Evidence from Five European Countries, *International Economic Journal*, 22:4, 445-459.

Fountas, S., M. Karanasos and A. Mendoza, 2004, Output Variability and Economic Growth: the Japanese Case. *Bulletin of Economic Research* 56, no. 4: 35363.

Fountas, S., M. Karanasos, and J. Kim, 2002, Inflation and Output Growth Uncertainty and Their Relationship with Inflation and Output Growth. *Economics Letters* 75, no. 3: 293301.

Fountas, S., Karanasos, M. and J. Kim, 2006, Inflation Uncertainty, Output Growth Uncertainty and Macroeconomic Performance. *Oxford Bulletin of Economics and Statistics* 68, 319-343.

Frank, J. and P. Garcia, 2009, Bid-Ask Spreads, Volume, and Volatility: Evidence from Livestock Markets, Selected Paper prepared for presentation at the Agricultural & Applied Economics Association 2009, AAEA & ACCI Joint Annual Meeting, Milwaukee, Wisconsin, July 26-29, 2009

Friedman, M., 1953, The Case for Flexible Exchange Rates, in *Essays in positive economics*, University of Chicago, Chicago.

Friedman, M., 1977, Nobel Lecture: Inflation and Unemployment. *Journal of Political Economy* 85, 451-472.

Frino, A. and S. Jones, 2005, The Impact of Mandated Cash Flow Disclosure on Bid-Ask Spreads. *Journal of Business Finance & Accounting* Vol. 32, No. 7-8, pp. 1373-1396.

French, K. R. and R. Roll, 1986, Stock Return Variances: The Arrival of Information and the Reactions of Traders. *Journal of Financial Economics* 17, 5526.

Fung, H. and G. Patterson, 1999, The Dynamic Relationship of Volatility, Volume, and Market Depth in Currency Futures Markets. *Journal of International Financial Markets, Institutions and Money* 9, 33-59.

Garman, M., and M. Klass, 1980, On the Estimation of Security Price Volatilities from Historical Data. *Journal of Business* 1, 67-78.

Ghysels, E., P. Santa-Clara, and R. Valkanov, 2005, There is a Risk-return Trade-off after All. *Journal of Financial Economics* 76, 509-548.

Gilbert, C. L. and H. Rijken, 2006, How is Futures Trading Affected By the Move to a Computerized Trading System? Lessons from the LIFFE FTSE 100 Contract. *Journal of Business Finance & Accounting* Vol. 33, No. 7-8, pp. 1267-1297.

Gillemot, L., J. D. Farmer and F. Lillo, 2005, There's More to Volatility than Volume. *physics* 0510007

Giot, P. and S. Laurent, 2003, Value-at-Risk for Long and Short Positions. *Journal of Applied Econometrics* 18, 641-663.

Glosten, L. and P. Milgrom, 1985, Bid, Ask and Transaction Prices in A Specialist Market with Heterogeneously Informed Traders. *Journal of Financial Economics* 14, 71–100.

Glosten, L. R. and L. E. Harris, 1988, Estimating the Components of the Bid-ask Spread, *Journal of Financial Economics* Vol. 21, 123-142.

Green, C. J., P. Maggioni and V. Murinde, 2000, Regulatory Lessons for Emerging Stock Markets from A Century of Evidence on Transactions Costs and Share Price Volatility in the London Stock Exchange. *Journal of Banking and Finance* 24 (4), pp. 577-601

Gregoriou, A., C. Ioannidis and L. Skerratt, 2005, Information Asymmetry and the Bid-Ask Spread: Evidence From the UK. *Journal of Business Finance&Accounting* 32(9)&(10), November/December

Gregoriou, A, J. Healy and C. Ioannidis, 2007, Market Frictions in the FTSE 100 Index Option Market: Evidence from A Data Mining Approach, *Journal of Futures Markets* Vol. 27, 471-494.

Grier, K., T. Henry, N. Olekalns and K. Shields, 2004, The Asymmetric Effects of Uncertainty on Inflation and Output Growth. *Journal of Applied Econometrics* 19, 551-565.

Grier, K. and G. Tullock, 1989, An Empirical Analysis of Cross-National Economic Growth, 1951-1980. *Journal of Monetary Economics* 24, 48-69.

Grier, K. and M. Perry, 1998, On Inflation and Inflation Uncertainty in the G7 Countries. *Journal of International Money and Finance* 17, 671-689.

Grier, K. and M. Perry, 2000, The Effects of Real and Nominal Uncertainty on Inflation and Output Growth: Some GARCH-M Evidence. *Journal of Applied Econometrics* 15, 45-58.

Grier, K, and G. Tullock, 1989, An Empirical Analysis of Cross-national Economic Growth: 1951-1980. *Journal of Monetary Economics* 24, no. 2: 25976.

Grinols, E., and S. J. Turnovsky, 1998, Risk, Optimal Government Finance, and Monetary Policies in a Growing Economy. *Economica* 65, no. 259: 40127.

Grundfest, J. and J. Shoven, 1991, Adverse Implications of a Securities Transactions Excise Tax. *Journal Account Audit Finance* 6:409–442

Haber, S. and E. Perotti, 2007, The Political Economy of Finance, Stanford and Amsterdam, mimeo.

Hamilton, J., 1989, A New Approach to the Economic Analysis of Nonstationary Time Series and the Business Cycle. *Econometrica* 57, no. 2: 35784.

Hau, H., 2006, The Role of Transaction Costs for Financial Volatility: Evidence from the Paris Bourse, *Journal of the European Economic Association*, 4(4), 862-890.

He, C. and T. Teräsvirta, 1999, Statistical Properties of the Asymmetric Power ARCH Model, in Engle, R. F. and H. White (eds), *Cointegration, Causality and Forecasting. Festchrift in honour of Clive W. J. Granger*. Oxford University Press, Oxford, 462-474.

Heflin, F, and K.W. Shaw, 2000, Blockholder Ownership and Market Liquidity, *Journal of Financial and Quantitative Analysis*, Vol. 35, 621-633.

Henry, O., and N. Olekalns, 2002, The Effect of Recessions on the Relationship between Output Variability and Growth. *Southern Economic Journal* 68, no. 1: 68392.

Hentschel, L., 1995, All in the Family. Nesting Symmetric and Asymmetric GARCH Models. *Journal of Financial Economics* 39, 71-104.

Holland, S., 1995, Inflation and Uncertainty: Tests for Temporal Ordering. *Journal of Money, Credit, and Banking* 27, 827-837.

Huang, R. D., and H. R. Stoll, 1997, The Components of the Bid-Ask Spread: A General Approach, *Review of Financial Studies* Vol. 10, 995-1034.

Jiranyakul, K. and T. P. Opiela, 2010, Inflation and Inflation Uncertainty in the ASEAN-5 Economies. *Journal of Asian Economics* 21, 105–112.

Jones, C. M., C. Kaul, M. L. Lipso, 1994, Information, Trading, and Volatility. *Journal of Financial Economics* 36, 127-154.

Jones, C. M. and P. J. Seguin, 1997, Transaction Costs and Price Volatility: Evidence from Commission Deregulation. *American Economic Review* 87(4), 728-737.

Kaminsky, G. and S. Schmukler, 2003, Short-Run Pain, Long-run Gain: The Effects of Financial Liberalization, NBER Working Paper No. 9787.

Kanagaretnam, K., G. Lobo and D. Whalen, 2005, Relationship between Analyst Forecast Properties and Equity Bid-Ask Spreads and Depths around Quarterly Earnings Announcements. *Journal of Business Finance & Accounting* 32(9)&(10),p1773-1799.

Karanasos, M., M. Karanassou and S. Fountas, 2004, Analyzing US Inflation by a GARCH Model with Simultaneous Feedback, *WSEAS Transactions on Information Science and Applications* 2, 767-772.

Karanasos, M., and A. Kartsaklas, 2009, Dual Long-memory, Structural Breaks and the Link between Turnover and the Range-based Volatility. *Journal of Empirical Finance* 16, 838-851.

Karanasos, M. and J. Kim, 2006, A Re-examination of the Asymmetric Power ARCH Model, *Journal of Empirical Finance* 13, 113-128.

Karanasos, M. and S. Schurer, 2005, Is the Reduction in Output Growth Related to the Increase in its Uncertainty? The Case of Italy. *WSEAS Transactions on Business and Economics* 3, 116-122.

Karanasos, M. and S. Schurer, 2008, Is the Relationship Between Inflation and its Uncertainty Linear? *German Economic Review* 9(3), 265-286.

Karanasos, M. and N. Zeng, 2007, UK Inflation, Output Growth and their Uncertainties: Four Variables, Twelve Links and Many Specifications. Unpublished Paper, Brunel University.

Karpoff, J. M., 1986, A Theory of Trading Volume. *Journal of Finance* Vol. 41, pp. 1069–87.

Kehoe, T. J., 2003, What Can We Learn from the Current Crisis in Argentina? *Scottish Journal of Political Economy* 50, 609-33.

Kehoe, T. J. and E. C. Prescott (Eds.), 2007, *Great Depressions of the Twentieth Century*, Federal Reserve Bank of Minneapolis.

Keynes, J. M., 1936, *The General Theory of Employment, Interest, and Money*. London: Macmillan.

Kim, O. and R. E. Verrecchia, 1994, Market Liquidity and Volume Around Earnings Announcements. *Journal of Accounting and Economics* Vol. 17, pp. 41–67.

King, R. G. and R. Levine, 1993, Finance and Growth: Schumpeter Might be Right. *Quarterly Journal of Economics* 108, 717-737.

King, R., C. Plosser and S. Rebelo, 1988, Production, Growth, and Business Cycles: II. New Directions. *Journal of Monetary Economics* 21, no. (2-3): 30941.

Kneller, R. and G. Young, 2001, Business Cycle Volatility, Uncertainty, and Long-run Growth. *Manchester School* 69, no. 5: 53452.

Kormendi, R. and P. Meguire, 1985, Macroeconomic Determinants of Growth: Cross-country Evidence. *Journal of Monetary Economics* 16, no. 2: 14163.

Kupiec, P., 1996, Noise Traders, Excess Volatility, and a Securities Transaction Tax. *Journal of Financ Serv Res* 10:115–129.

Kydland, F. and E. Prescott, 1982, Time to Build and Aggregate Fluctuations. *Econometrica* 50, no. 6: 134570.

Laurent, S., 2004, Analytical Derivatives of the APARCH Model. *Computational Economics* 24, 51-57.

Lee, C. M., C. B. Mucklow. and M. J. Ready, 1993, Spreads, Depths and the Impact of Earnings Information: An Intraday Analysis. *Review of Financial Studies* Vol 6, 345-374.

Levine, R., 2005, Finance and Growth: Theory and Evidence. In Aghion, P. and S. Durlauf (Eds.), *Handbook of Economic Growth*. Elsevier, Amsterdam.

Lin, J, G. Sanger, and G. G. Booth, 1995, Trade Size and Components of the Bid-Ask Spread, *Review of Financial Studies* Vol. 8, 1153-1183.

Liu, S., Z. Zhu, 2009, Transaction Costs and Price Volatility: New Evidence from the Tokyo Stock Exchange. *Journal of Financial Services Research* 36 (1), pp. 65-83.

Loayza, N. V. and R. Rancière, 2006, Financial Development, Financial Fragility and Growth. *Journal of Money Credit and Banking* 38, 1051-1076.

Long, J. and C. Plosser, 1983, Real Business Cycles. *Journal of Political Economy* 91, no. 1: 3969.

Lobato, I. and C. Velasco, 2000, Long Memory in Stock-market Trading Volume. *Journal of Business & Economic Statistics* 18, 410-427.

Lynch, J., 1985, Argentina: From Independence to National Organization, in Leslie Bethell (ed) *The Cambridge History of Latin America* Volume 3, From Independence to c.1870, The Cambridge History of Latin America

Madhavan, A., M. Richardson and M. Roomans, 1997, Why do Security Prices Change? A transaction-level analysis of NYSE stocks. *Review of Financial Studies* Vol. 10, 1035-1064.

Maddison, A., 2007. *Contours of the World Economy* I-2030 AD. Oxford University Press, Oxford.

McConnell, M., and G. Perez-Quiros, 2000, Output Fluctuations in the United States: What has Changed since the Early 1980s? *American Economic Review* 90, no. 5: 146476.

McGroartya, F., O. Gwilymb, S. Thomasc, 2009, The Role of Private Information in Return Volatility, Bid–ask Spreads and Price Levels in the Foreign Exchange Market. *Journal of International Financial Markets, Institutions & Money* 19, 387–401.

Menyah, K. and K, Paudyal, 2000, The Components of Bid-ask Spreads on the London Stock Exchange. *Journal of Banking and Finance* Vol. 24, 1767-1785

Miller, M. H., 1991, *Financial innovations and market volatility*, Oxford: Basil Blackwell.

Mirman, L., 1971, Uncertainty and Optimal Consumption Decisions. *Econometrica* 39, no. 1: 17985.

Mitchell, R., 1998, *International Historical Statistics: Europe*. Macmillan.

Mitchell, R., 2003, *International Historical Statistics: The Americas*, 1750-2000. Palgrave MacMillan, London.

Mittnik, S. and M. Paolella, 2000, Conditional Density and Value-at-Risk Prediction of Asian Currency Exchange Rates. *Journal of Forecasting* 19, 313-333.

Mulherin, H. J., 1990, Regulation, Trade Volume, and Stock Market Volatility. *Revue Economique*, 41(5), 923-937.

Murdoch, J. and T. Sandler, 2004, Civil Wars and Economic Growth: Spatial Dispersion. *American Journal of Political Science* 48, 138-151.

Norden, L. L., 2009, A Brighter Future with Lower Transactions Costs? *The Journal of Futures Markets* Vol. 29, No. 8, 775-796

Palm, F. C., 1996, GARCH Models of Volatility. in: Maddala, G. S., Rao, C. R. (eds.), *Handbook of Statistics: Statistical Methods in Finance*, Vol. 14 North Holland, Amsterdam, 209-240.

della Paolera, G. and A. M. Taylor, 1998, Finance and Development in an Emerging Market: Argentina in the Interwar Period. In: *Latin America and the World Economy since 1800*, 139-169. Harvard University Press, Cambridge.

della Paolera, G. and A. M. Taylor, 2003 (eds). *A New Economic History of Argentina*. Cambridge University Press.

Pesaran, H., 1997, The Role of Econometric Theory in Modeling the Long Run. *Economic Journal* 107, 178-191.

Pesaran, H. and Y. Shin, 1999, An Autoregressive Distributed Lag Modeling Approach to Cointegration. in *Econometrics and Economic Theory in the 20th Century: The Ragnar Frisch Centennial Symposium*, chap 4, pp. 371-413. Cambridge University Press.

Pindyck, R., 1991, Irreversibility, Uncertainty, and Investment. *Journal of Economic Literature* 29, no. 3: 111048.

Pinto de Andrade, A. and J. A. Divino, 2005, Monetary policy of the Bank of Japan—inflation target versus exchange rate target. *Japan and the World Economy* 17 189–208.

Pirrong, C., 1996, Market Liquidity and Depth on Computerized and Open Outcry Trading Systems: A Comparison of the DTB and LIFFE Bund Contracts. *Journal of Futures Markets* 16(5):519-543.

Prados de la Escosura, L. and I. Sanz-Villarroya, 2009, Contract Enforcement, Capital Accumulation and Argentina's Long-run Decline. *Cliometrica* 3, 1–26.

Prasad, E. S., Rogoff, K. S., S. Wei. and M. A. Kose, 2004, Financial Globalization, Growth and Volatility in Developing Countries. NBER Working Paper 10942.

Puffer, M., 1991, Private information and weekend volatility in the Tokyo and New York stock markets. *Journal of Banking and Finance* I5, 407-423. North-Holland

Ramey, G. and V. Ramey, 1991, Technology Commitment and the Cost of Economic Fluctuations. NBERWorking Paper No. 3755.

Ramey, G. and V. Ramey, 1995, Cross-country Evidence on the Link between Volatility and Growth. *American Economic Review* 85, 1138-1151.

Rock, D., 1985, Argentina in 1914: The Pampas, The Interior, Buenos Aires, in Leslie Bethell (ed) *The Cambridge History of Latin America*, Cambridge University Press, Cambridge.

Roll, R., 1984, A Simple Implicit Measure of the Effective Bid-Ask Spread in an Efficient Market. *Journal of Finance* Vol. 39, pp. 1127–40.

Roll, R., 1989, Price Volatility, International Market Links, and Their Implications for Regulatory Policies. *J Financ Serv Res* 3:211–246.

Romer, C., 2007, *Great depression. In Encyclopaedia Britannica*. Encyclopaedia Britannica, Incorporated.

Sandmo, A., 1970, The Effect of Uncertainty on Saving Decisions. *Review of Economic Studies* 37, no. 3: 35360.

Sanz-Villarroya, I., 2005, The Convergence Process of Argentina with Australia and Canada: 1875–2000. *Explorations in Economic History* 42, 439-458.

Sanz-Villarroya, I., 2007, La Belle Époque de la Economía Argentina. 1875-1913. *Acciones e Investigaciones Sociales* 23, 115-138.

Saporta, V. and K. Kan, 1997, The Effects of Stamp Duty on the Level and Volatility of UK Equity Prices, Working Paper, Bank of England.

Schwert, G. W., 1990. Stock Volatility and Crash of ’87. *Review of Financial Studies* 3 (1), 77–102. Sims, C.A., 1980. Macroeconomics and reality. Econometrica 48 (1), 1–48.

Shen, C. H. and D. Hakes, 1995, Monetary Policy as a Decision-Making Hierarchy: The Case of Taiwan. *Journal of Macroeconomics* 17, 357–368.

Smith, W. T., 1996, Taxes, Uncertainty and Long-term Growth. *European Economic Review* 40, no. 8: 164764.

Solberg, C. E., 1987, *The Prairies and the Pampas: Agrarian Policy in Canada and Argentina: 1880-1930*. Stanford University Press, Stanford.

Speight, A., 1999, UK Output Variability and Growth: Some Further Evidence. *Scottish Journal of Political Economy* 46, no. 2: 17584.

Stiglitz, J., 1989, Using Tax Policy to Curb Speculative Short-term Trading. *J Financ Serv Res* 3:101–115.

Stock, J. and M. Watson, 2002, Has the Business Cycle Changed and Why? National Bureau of EconomicResearch Working Paper no. 9127, September.

Stoll, H. R., 1978, The Pricing of Security Dealer Services: An Empirical Study of NASDAQ Stocks. *Journal of Finance* 33, 1153–1172.

Stoll, H. R., 1989, Inferring the Components of the Bid-ask Spread: Theory and Empirical Tests. *Journal of Finance* 44, 115-134.

Summers, L., V. Summers, 1989, When Financial Markets Work too Well: A Cautious Case for a Securities Transaction Tax. *J Financ Serv Res* 3:261–286.

Taylor, A., 1992, External Dependence, Demographic Burdens and Argentine Economic Decline after the Belle Epoque. The Journal of Economic History 52(4), 907-936.

Taylor, A., 2003. Capital Formation: Saving, Investment, and Foreign Capital. In della Paolera, G. and Taylor, A. (Eds.), *A New Economic History of Argentina*.

Thornton, J., 2007, The Relationship between Inflation and Inflation Uncertainty in Emerging Market Economies. *Southern Economic Journal* 73(4), 858-870.

Tobin, J., 1978, A Proposal for International Monetary Reform. *Eastern Economic Journal* 4, 153-159.

Tobin, J., 1984, On the Efficiency of the Financial System, *Lloyds Bank Review* 153 (July), 1-15.

Tornell, A., F. Westermann and L. Martinez, 2004, The Positive Link Between Financial Liberalization, Growth and Crises, NBER Working Paper No. 10293.

Tsay, W. and C. Chung, 2000, The Spurious Regression of Fractionally Integrated Processes. *Journal of Econometrics* 96, 155-182.

Umlauf, S. R., 1993, Transaction Taxes and the Behavior of the Swedish Stock Market, *Journal of Financial Economics* 33, 227-240.

Ungar, M. and B. Zilberfarb, 1993, Inflation and its Unpredictability - Theory and Empirical Evidence. *Journal of Money, Credit, and Banking* 25, 709-720.

Van Ness, B. F., R. A. Van Ness and R. S. Warr, 2001, How Well Do Adverse Selection Components Measure Adverse Selection? *Financial Management* Vol. 30, 77-98.

Véanzonès, M. and C. Winograd, 1997, *Argentina in the 20th Century. An Account of Long-Awaited Growth*. Organization for Economic Co-operation and Development, Paris.

Visco, I., 1999, Structural Reform in Korea after the 1997 Economic Crisis, the Agenda and Implementation. Paper presented in the Conference on "Economic Crisis and Restructuring in Korea", Seoul, 3rd December.

Wang, G. H. K. and J. Yau, 2000, Trading Volume, Bid-Ask Spread, and Price Volatility in Futures Markets. *Journal of Futures Markets* 20(10):943-970.

Wiggins, H., 1992, Estimating the Volatility of S&P 500 Futures Prices Using the Extreme-value Method. *Journal of Future Markets* 12, 265-273.